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Does austerity pay off?

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Abstract

We investigate if a reduction of government consumption lowers the sovereign default premium. For this purpose we build a new data set for 38 emerging and developed economies. Results vary along three dimensions. First, the time horizon: the premium declines, but only in the long run. Second, initial conditions: the premium increases in the short run, but only if it is already high. Third, size: the short-run response of the premium increases disproportionately as government consumption is reduced. We rationalize these findings in a structural model of optimal sovereign default where default risk is priced in an actuarially fair manner.

Keywords*: Fiscal policy, austerity, sovereign risk, default premium, local projections, panel VAR, fiscal stress*

Jel codes: E62, E43, C32

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1 Introduction

In the years following the global financial crisis, many European governments implemented sizeable austerity measures. These included spending cuts and tax increases and were meant to confront concerns about rising levels of public debt or outright solvency issues. In fact, yields on debt issued by several European sovereigns started to take off by 2010, reflecting sizeable default premiums (Krishnamurthy et al., 2014). Still, as austerity measures were implemented, default premiums kept rising. Moreover, as Figure 1 illustrates, large spending cuts were associated with strong increases of the default premium. Against this background, we ask whether austerity actually pays off. Specifically, we ask whether austerity causes the default premium to decline and, if so, when and under which circumstances.

We focus on how financial markets respond to austerity measures and sidestep the issue of how such measures impact the actual health of government finances. While austerity impacts fiscal fundamentals such as the level of sovereign debt, these fundamentals typically fail to provide a sufficient statistic for assessing the sustainability of debt. For the ability and willingness of governments to service its debt obligations and to roll over liabilities also depends on market conditions (Calvo, 1988; Cole and Kehoe, 2000; Roch and Uhlig, 2015) and a number of country-specific, partly unobserved factors such as the ability to raise tax revenues (Bi, 2012; Lorenzoni and Werning, 2014; Trabandt and Uhlig, 2011). The same level of debt may thus have very different implications for debt sustainability at different times and in different countries. The default premium demanded by financial markets, instead, is based on a broader assessment and provides a more comprehensive picture.

To take up the issue, we assemble a new data set for the sovereign default premium in a large number of countries, including countries outside of the euro area. Specifically, we construct time series for the sovereign default premium in 38 developed and emerging economies, covering the period since the early 1990s. We compute the default premium as the difference in sovereign yields vis-à-vis a riskless reference country where sovereign default can be ruled out for practical purposes. Importantly, we only consider yields on

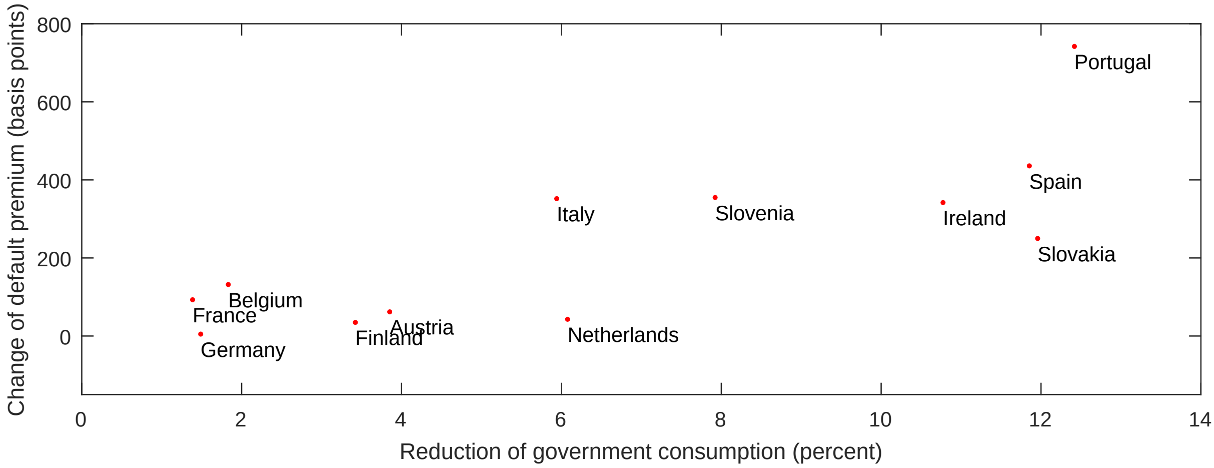


Figure 1: Sovereign default premium and austerity in selected euro area economies: 2010Q1–2012Q2. Notes: vertical axis measures change of default premium in basis points, horizontal axis measures reduction of real government consumption in percent, see Section 2 for detailed description of the data.

government securities issued in a common currency in order to eliminate confounding factors due to expectations of inflation and currency depreciation. For more recent observations, we also rely on data for credit default swap (CDS) spreads. We analyze the variation of the default premium in some detail, both across countries and over time, and summarize it by the cumulative distribution function.

Our interest is on how the sovereign default premium responds to austerity measures. Because we lack high quality data on taxes and transfers, we investigate how changes of government consumption impact the default premium.[1] In doing so, we focus on three dimensions. First, we distinguish between the response of the default premium in the short run and in the long run. In this regard, we obtain an important result: austerity pays off, but only in the long run. Specifically, we find that a cut of government consumption by one percent reduces the default premium by about 17 basis points in the long run. In contrast, during the first two years after the cut, the premium increases sharply. At the same time, we find a significant decline of output.

Second, we use the cumulative distribution function of the sovereign default premium to classify initial conditions in terms of "fiscal stress": the premium rises in the short run only if the premium is high to begin with. This condition is often met when austerity

[1]Some authors find that whether austerity is tax based or spending based is crucial for how it impacts the economy (see, for recent contributions, Alesina and Ardagna, 2013; Alesina et al., 2015a,b).

measures are implemented, as we document in ongoing work (Born et al., 2016). Third, we vary the size of the shock. While cutting government consumption pushes up the premium, the converse does not apply: an increase of government consumption in times of fiscal stress leaves the default premium basically unchanged. More generally, we document that size matters: large cuts of government consumption have a disproportionate effect on the premium in the short run—provided fiscal stress is high. The differences along these dimensions turn out to be statistically significant. They also matter economically, as they easily amount to some 50 basis points, given a change of government consumption by one percent.

We use a number of alternative time-series models to establish these results.[2] In terms of identification we draw on earlier work by Blanchard and Perotti (2002) and Ramey (2011b). Importantly, we assume that government consumption is predetermined relative to output and the sovereign default premium: we interpret any variation of government consumption which is unaccounted for by past observations as a structural innovation. We dispel two major concerns which may arise in this context. First, policy makers may respond swiftly and systematically to increases of the default premium, notably in times of fiscal stress. We show that this is not the case by identifying a common factor of the sovereign default premium in the cross-section of our panel data set. While government consumption is indeed reduced in response to a higher (common) default premium, it takes time for this effect to materialize. Second, market participants may anticipate fiscal measures. To address this concern, we purge our measure of fiscal shocks by the forecast of government consumption, which is available for a subset of our sample.

Our findings for the short run seemingly support the notion that financial markets are "schizophrenic" about austerity in that they demand austerity measures as public debt builds up, but fail to reward them as austerity slows down output growth (Blanchard, 2011; Cotarelli and Jaramillo, 2013). To investigate this issue, we rely on a structural model of optimal sovereign default. Specifically, we modify the model of Arellano (2008) by

[2] We build on the classic studies of the fiscal transmission mechanism within vector autoregressive models (Blanchard and Perotti, 2002; Mountford and Uhlig, 2009; Ramey, 2011b), but also on local projections (Jordá, 2005). Throughout, we allow the effects of austerity measures to be non-linear.

allowing for exogenous variation in government consumption and a multiplier effect of such variation on output. The model predicts, in line with the evidence, that cutting government consumption temporarily raises the default premium if fiscal stress is high. This, however, is not because financial market participants are schizophrenic about austerity. Rather, this is because, as long as output is temporarily reduced due to austerity, investors understand that a government is tempted to default on its debt obligations in order to free scarce resources from debt service for other expenditures.

Our paper relates to a number of empirical studies which explore the state-dependence of the effects of fiscal policy (e.g., Auerbach and Gorodnichenko, 2012; Ramey and Zubairy, 2014). While the issue remains controversial to date, earlier work by Perotti (1999) established that fiscal policy affects the economy differently in “good times” and “bad”. More recently, Corsetti et al. (2012a), Auerbach and Gorodnichenko (2013), and Ilzetzki et al. (2013) find that the fiscal multiplier depends on the level of public debt. Bertola and Drazen (1993) and Corsetti et al. (2013) provide model-based analyses. Giavazzi et al. (2000) suggest that the size and persistence of fiscal measures also matters for their effects. Earlier studies have also focused on how financial markets respond to fiscal policy measures. Ardagna (2009), for instance, reports that interest rates tend to decline in response to large fiscal consolidations. Laubach (2009) finds that future debt and deficits tend to raise U.S. interest rates. Akitoby and Stratmann (2008) focus on how sovereign yield spreads in emerging markets react to changes of fiscal indicators.

The remainder of the paper is organized as follows. Section 2 details the construction of our data set. In this section, we also establish a number of basic facts regarding the time-series properties of the sovereign default premium and its relationship to government consumption and output growth. We discuss our econometric approach and our results in Sections 3 and 4. Section 5 provides a structural interpretation through the lens of a model of optimal sovereign default. Section 6 offers some conclusions.

2 Data

Our analysis is based on a new data set. It contains quarterly observations for government consumption, output, and the sovereign default premium in 38 emerging and advanced economies. While data on default premiums are available at higher frequency, data on macroeconomic aggregates are not. For a long time, time-series studies of the fiscal transmission mechanism have been limited to a small set of countries because high-quality quarterly data for government consumption was not available.[3] Rather, quarterly data was often derived from indirect sources using time disaggregation/interpolation. In a recent contribution, Ilzetzki et al. (2013) have collected quarterly data based on direct sources for government consumption for 44 countries. Quarterly data of a comparable coverage for other fiscal variables such as taxes, transfers, or deficits are not available; hence our focus on government consumption.

We collect quarterly data for government consumption expenditure based on national accounts/non-financial accounts of the government along the lines of Ilzetzki et al. (2013). On the one hand, we limit our focus to those countries for which we are also able to compute a sovereign default premium. On the other hand, we extend their sample to include more recent observations and additional countries for which we were able to confirm with statistical agencies the availability of government consumption data based on direct sources.[4] The full sample coverage is shown in Table 1. Our earliest observation for which we obtain data on the default premium and on government consumption is 1991Q1, namely for Denmark and Italy. Our sample runs up to 2014.

Table 1 also provides summary statistics for the government consumption-to-GDP ratio for our sample where both government consumption data and default premiums are available. Government consumption from national accounts/non-financial accounts of the

[3]Some studies have resorted to annual data (e.g., Beetsma et al., 2006, 2008; Bénétrix and Lane, 2013). In this case identification assumptions tend to be more restrictive. However, Born and Müller (2012) consider both quarterly and annual data for four OECD countries. They find that the estimated effects of government spending shocks do hardly differ.

[4]For several European countries, we also include earlier observations for the 1990s whenever we are able to compute a default premium. In this case, government consumption data is available through Eurostat. However, it is not entirely based on direct sources, implying that the data falls short of the more recent Eurostat standards, firmly established for data since the mid-2000s only. We therefore verify below that our results are robust with respect to employing a more conservative sample.

Table 1: Basic properties of government consumption-to-GDP ratio

Country	first obs	last obs	min	max	mean	std
Argentina	1994Q1	2013Q3	11.8	18.3	13.9	1.6
Australia	2003Q2	2010Q3	17.0	18.1	17.5	0.3
Austria	1994Q1	2014Q1	17.8	20.5	19.1	0.7
Belgium	1995Q1	2014Q1	21.1	25.3	22.8	1.4
Brazil	1995Q1	2014Q1	18.8	22.6	20.4	0.8
Bulgaria	1999Q1	2014Q1	14.1	20.5	17.6	1.5
Chile	1999Q3	2014Q2	4.5	6.5	5.6	0.5
Colombia	2000Q1	2014Q1	15.1	17.3	16.3	0.5
Croatia	2004Q2	2014Q1	18.2	20.6	19.5	0.6
Czech Republic	2004Q2	2014Q1	19.3	22.0	20.8	0.7
Denmark	1991Q1	2014Q1	24.7	30.1	26.4	1.5
Ecuador	1995Q2	2014Q1	8.9	14.2	11.6	1.5
El Salvador	2002Q3	2014Q1	6.4	8.8	7.3	0.6
Finland	1992Q3	2014Q1	19.6	25.0	22.1	1.6
France	1999Q2	2014Q1	22.7	25.1	23.8	0.7
Germany	2004Q2	2014Q1	17.7	20.1	19.0	0.6
Greece	2000Q1	2011Q1	16.9	22.3	18.0	1.1
Hungary	1999Q2	2014Q1	20.2	24.7	21.8	1.1
Ireland	1997Q1	2014Q1	13.9	20.8	16.7	1.7
Italy	1991Q1	2014Q1	17.5	21.6	19.4	1.0
Latvia	2006Q2	2014Q1	15.5	21.8	18.0	1.6
Lithuania	2005Q3	2014Q1	16.6	22.1	18.8	1.6
Malaysia	2000Q1	2014Q1	6.6	12.1	9.8	1.2
Mexico	1994Q1	2014Q2	0.3	0.5	0.4	0.1
Netherlands	1999Q2	2014Q2	20.2	26.8	23.8	2.2
Peru	1997Q2	2014Q2	6.9	9.4	8.0	0.5
Poland	1995Q1	2014Q1	16.8	19.9	18.1	0.6
Portugal	1995Q1	2014Q1	17.2	22.3	19.5	1.4
Slovakia	2004Q2	2014Q1	16.8	20.1	18.3	0.9
Slovenia	2003Q2	2014Q1	17.1	21.0	19.5	1.2
South Africa	1995Q1	2014Q1	17.9	22.8	19.6	1.4
Spain	1995Q1	2014Q1	16.9	21.8	18.6	1.6
Sweden	1993Q2	2014Q2	6.7	9.8	7.8	0.8
Thailand	1997Q3	2014Q2	9.8	14.4	12.1	1.2
Turkey	1998Q1	2014Q1	9.7	15.7	12.8	1.4
United Kingdom	1993Q1	2013Q4	17.5	23.4	20.2	1.6
United States	2008Q1	2014Q1	14.7	17.1	16.1	0.7
Uruguay	2001Q3	2014Q1	10.0	14.2	11.1	1.2

Notes: Government consumption is consumption of the general government except for Chile, El Salvador, Malaysia, Mexico, Peru, and Sweden, where it refers to central government consumption. The government consumption-to-GDP ratio is computed as the ratio of nominal variables, except for Uruguay, where we compute it as the ratio of real variables. The ratio is measured in percentage points. For Mexico, the share of central government wages and goods and services purchases is only a very small share of GDP.

government is exhaustive government final consumption. It is accrual based and does not include transfer payments or government investment (see Lequiller and Blades, 2006, Chapter 9). Depending on the availability of quarterly time series, it pertains to either the general or the central government. The ratio of government consumption-to-GDP varies both across time and across countries. In case of general government data, government consumption fluctuates around 18 percent of GDP.

As a distinct contribution, we also construct a panel data set for the sovereign default premium in order to measure the assessment of financial markets regarding the sustainability of public finances. Given observations on quarterly government consumption, we aim to construct measures of default risk for as many countries as possible. As stressed in the introduction, we construct a mostly spread-based measure using yields for securities issued in common currency. To the extent that goods and financial markets are sufficiently integrated, we thus eliminate fluctuations in yields due to changes in real interest rates, inflation expectations, and the risk premiums associated with them. In addition to a default risk premium, if duration differs or drifts, yield spreads may still reflect a term premium (see Broner et al., 2013). We try to minimize the term premium by constructing the yield spread on the basis of yields for bonds with a comparable maturity and coupon.[5] As a result, yield spreads should primarily reflect financial markets' assessment of the probability and extent of debt repudiation by a sovereign.[6]

We obtain our default-risk measure based on four distinct sources/strategies. First, for a subset of (formerly) emerging markets we directly rely on J.P. Morgan's Emerging Markets Bond Index (EMBI) spreads, which measure the difference in yields between dollar-denominated government or government-guaranteed bonds of a country and U.S. government bonds.[7]

Second, we add to those observations data for euro area countries based on the "long-term interest rate for convergence purposes". Those are computed as yields to maturity from "long-term government bonds or comparable securities" with a residual maturity of close to 10 years with sufficient liquidity (for details, see European Central Bank,

[5]We focus on long-term rates whenever possible. As they are closely linked to the average of expected future short-term rates, they are a more appropriate measure of governments' refinancing costs than short-term rates. Assessing the effects of austerity on the term structure is beyond the scope of the present study.

[6]In principle, spreads may also reflect a liquidity premium—an issue we ignore in what follows because we consider government debt traded in mature markets. See Appendix A.1.3 for a more detailed discussion.

[7]See Appendix A.1.1 for details on the EMBI.

2004). For this country group, we use the German government bond yield as the risk-free benchmark rate and compute spreads relative to the German rate.[8]

Third, we make use of the issuance of foreign-currency government bonds in many advanced economies during the 1990s and 2000s to extend our sample to non-euro area countries and the pre-euro period. In case of countries like Denmark, Sweden, or the UK, this allows us to compute common-currency yield spreads, even though those countries are not members of the euro area. Drawing on earlier work by Bernoth et al. (2012), we identify bonds denominated in either U.S. dollar or Deutsche mark of at least 5 years of maturity issued by developed economies. We compute the yield spread for those bonds relative to the yields of U.S. or German government bonds of comparable maturity and coupon yield.[9] Whenever possible, we aim to minimize the difference in coupon yield to 25 basis points and the difference in maturity to one year. In order to avoid artifacts because trading dries up in the last days before redemption, we omit the last 30 trading days before the earliest maturity date of either the benchmark or the government bond.[10] In case that several bonds are available for overlapping periods, we average over yield spreads using the geometric mean. This procedure mimics the creation of the EMBI spreads and the "long-term interest rate for convergence purposes". However, we necessarily rely on a smaller foreign currency bond universe and cannot correct for maturity drift. Thus, we rely on the "long-term interest rate for convergence purposes" whenever they are available.[11]

[8]The bonds used for computing the "long-term interest rate for convergence purposes" are typically bonds issued in euro, but under national law. In this regard they differ from the securities on which the EMBI is based, which are typically issued under international law. This difference becomes important if the monetary union is believed to be reversible. In case of exit from the EMU, the euro bonds will most likely be converted into domestic-currency bonds, implying that they should carry a redenomination premium that is absent in case of international-law bonds. Still, even during the height of the European debt crisis, the redenomination premium accounted for a moderate fraction of sovereign yield spreads (Krishnamurthy et al., 2014; Kriwoluzky et al., 2015). In any case, our main results also hold for a sample of emerging market countries.

[9]Yields on individual bonds are based on the yield to maturity at the midpoint as reported in Bloomberg or the yield to redemption in Datastream.

[10]Still, in moving along the yield curve, we may pick up cross-country differences in the slope of the yield curve. In principle, this effect can be quantitatively significant (Broner et al., 2013). However, as we find our spread measure to co-move very strongly with CDS spreads (whenever they are available), we ignore the issue in the present paper.

[11]By focusing on common-currency bonds, our spread measure is not affected by the convergence play observed for nominal yield spreads prior to the introduction of the euro.

Finally, in the more recent part of the sample, a direct measure of default risk has become available in the form of CDS spreads. Credit default swaps are insurance contracts that cover the repayment risk of an underlying bond. The CDS spread indicates the annual insurance premium to be paid by the buyer.[12] Accordingly, a higher perceived default probability on the underlying bond implies, ceteris paribus, a higher CDS spread. While well-suited to capture market assessment of debt sustainability, CDS spread data are generally only available after 2003 (see Mengle, 2007). Unfortunately, trading in these markets was often thin before the financial crisis, price discovery often took place in bond markets, and CDS contracts are subject to counterparty risk (see Fontana and Scheicher, 2010). Thus, we use CDS spreads to measure default risk only when no spread-based default premium measure is available.[13]

The use of CDS spreads also allows us to include the benchmark countries United States (EMBI) and Germany (long-term convergence yields) in the sample. In order to get an absolute measure of default risk for the other countries, we add the CDS spread of the respective benchmark countries to the relative country spread. For the period before CDS data are available, we add the value of the average CDS spread of the period prior to the default of Lehman Brothers.[14] Appendix A.3 provides an example of the construction of our data set.

Table 2 provides basic descriptive statistics for our absolute default-premium measure. The default premium s_t is measured in percentage points and varies considerably across our sample from which we exclude periods of default (see the table note). In a couple of euro-area countries the lowest realizations of the default premium are slightly negative.[15] For the group of developed economies (see Table 2 for the classification), we observe the highest premia in Portugal (12 pps) and Greece (10 pps). For emerging economies, the

[12]A no-arbitrage argument implies that the CDS spread should equal the spread between a par floating rate bond and the risk-free rate (Duffie, 1999).

[13]The CDS data construction is described in Appendix A.1.2. The correlation between CDS spreads and the yield-based default premium measures, when both are available, is typically above 0.9.

[14]Before the Lehman Brothers default, German and U.S. CDS were below 8 basis points and thus virtually zero. After Lehman, they peak at about 70 basis points and slowly return to about 15 basis points.

[15]The reason is that the long-term convergence yields are sometimes slightly lower than the German ones. This is presumably due to their construction not controlling for different bond duration characteristics and small maturity differences.

Table 2: Basic properties of sovereign default premia

Country	Group	min	max	mean	std	$\rho(\Delta y_t, s_t)$	$\rho(\Delta g_t, s_t)$
Argentina	E	2.12	19.50	7.78	3.65	-0.55	-0.06
Australia	D	0.03	1.30	0.31	0.31	-0.38	-0.39
Austria	D	0.03	1.98	0.40	0.41	-0.47	-0.31
Belgium	D	0.03	2.94	0.59	0.59	-0.42	-0.20
Brazil	E	1.64	24.20	5.70	4.17	-0.05	-0.07
Bulgaria	E	0.73	9.18	3.24	2.53	-0.11	-0.04
Chile	E	0.62	4.04	1.60	0.63	-0.46	0.10
Colombia	E	1.26	10.73	3.56	2.07	-0.40	-0.17
Croatia	E	0.15	5.47	2.07	1.61	-0.66	-0.47
Czech Republic	D	0.05	2.08	0.62	0.54	-0.83	-0.05
Denmark	D	0.02	2.18	0.53	0.46	-0.20	-0.05
Ecuador	E	5.09	21.20	9.86	4.07	-0.44	-0.36
El Salvador	E	1.34	9.15	3.56	1.45	-0.75	0.04
Finland	D	-0.02	1.27	0.39	0.29	-0.50	-0.15
France	D	0.03	2.03	0.44	0.46	-0.41	0.01
Germany	D	0.02	0.73	0.20	0.18	-0.34	0.07
Greece	D	0.18	10.02	1.49	2.58	-0.61	-0.21
Hungary	E	0.17	6.37	2.00	1.75	-0.60	-0.05
Ireland	D	-0.02	9.09	1.41	2.15	-0.19	-0.39
Italy	D	-0.03	5.86	0.98	1.18	-0.42	-0.39
Latvia	D	0.05	10.01	2.75	2.30	-0.72	-0.74
Lithuania	D	0.06	7.25	2.32	1.83	-0.65	-0.23
Malaysia	E	0.74	4.31	1.70	0.71	-0.65	-0.05
Mexico	E	1.18	15.96	3.65	2.58	-0.28	-0.04
Netherlands	D	-0.01	1.18	0.34	0.32	-0.63	-0.28
Peru	E	1.24	9.18	3.52	1.93	-0.26	0.02
Poland	E	0.49	8.78	2.02	1.33	-0.05	-0.12
Portugal	D	0.03	12.28	1.63	2.84	-0.45	-0.42
Slovakia	D	0.04	4.10	1.22	1.21	-0.39	-0.23
Slovenia	D	-0.15	5.42	1.58	1.78	-0.47	-0.44
South Africa	E	0.77	6.59	2.42	1.25	-0.54	-0.18
Spain	D	-0.03	5.40	0.95	1.35	-0.65	-0.45
Sweden	D	0.01	1.20	0.39	0.24	-0.33	-0.07
Thailand	E	0.27	5.62	1.38	0.92	-0.38	0.13
Turkey	E	1.89	10.73	4.48	2.40	-0.33	-0.16
United Kingdom	D	0.05	1.20	0.45	0.24	-0.43	-0.06
United States	D	0.07	0.61	0.27	0.12	-0.48	0.12
Uruguay	E	1.51	16.50	4.02	3.13	-0.42	-0.38

Notes: Default premium s_t is end-of-quarter observation, measured in percentage points. The last two columns report the correlation of default premiums with the growth rates of real GDP, Δy_t, and government consumption, Δg_t, respectively. Following IMF (2015, Tables B to D), group entry "D" denotes developed economies, while "E" denotes emerging economies. Excludes default episodes in Argentina (2001Q4–2005Q2), Ecuador (1999Q3–2000Q3 and 2008Q4–2009Q2), and Greece (2012Q1–2012Q2, 2012Q4) as classified by Standard & Poor's (see Chambers and Gurwitz, 2014, Table 2).

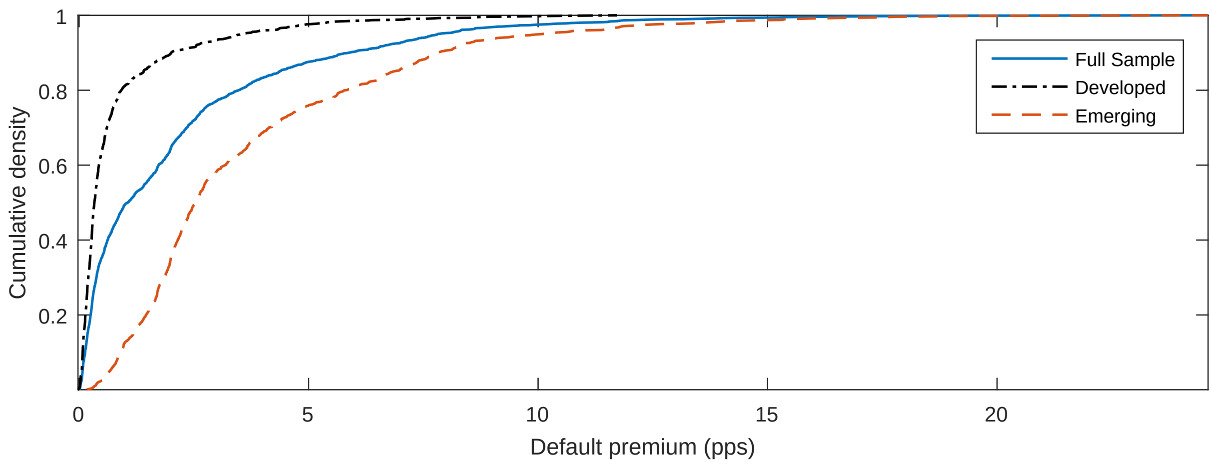


Figure 2: Sovereign default premium: empirical cumulative density function (CDF). Notes: horizontal axis measures the default premium in percentage points. Vertical axis measures fraction of observations for which the default premium is at most the value on the horizontal axis. Solid line displays CDF for full sample, dashed-dotted line: developed economies only, dashed line: emerging economies only.

highest values are reached in Brazil (24 pps), Ecuador (21 pps), and Argentina (20 pps).[16]

Compared to these values, most realizations of the default premium in our sample are small. This is apparent from the CDF plotted in Figure 2 for the entire sample (solid line), but also for the set of developed (dash-dotted line) and emerging economies (dashed line) in isolation. The total number of observations in our sample is 2320, of which 1247 are for developed economies and 1073 for emerging economies. In each case, the mass of observations is very much concentrated on the left. For the full sample about 50 percent of the observations for the default premium are below 1 percentage point. Still, there are considerable differences across the two country groups: 99.8 percent of observations are below 10 percentage points in the sample of developed economies. The corresponding number is only 95 percent in the sample of emerging-market economies.

Finally, in the last two columns of Table 2 we report the correlation of the sovereign default premium with output growth and the growth of government consumption, respectively. It turns out that the default premium is countercyclical in all countries, although sometimes the correlation is negligible. In contrast, the within-country correlation of

[16] During default episodes, spreads in secondary markets can achieve even higher values. In case of Argentina, the peak spread was 70 percentage points. Greek spreads were also higher shortly before and during the defaults (2012Q1–2012Q2, 2012Q4), but these observations are not included in our sample, because we lack high-quality national accounts data.

the default premium and government consumption growth varies across countries. It is negative for most of the countries, but often weakly so.

Eventually, we seek to establish the co-movement of the default premium and government consumption conditional on an exogenous variation of government consumption. In order to do so, we rely on specific identification assumptions which we discuss in what follows.

3 Evidence from vector autoregressions

We now specify and estimate a vector autoregression (VAR) model as a first step towards establishing the dynamic effects of austerity on the sovereign default premium. In section 4 we provide additional evidence based on local projections which offer, in some respects, more flexibility than VARs. Yet, because VAR models are frequently employed to characterize the fiscal transmission mechanism (Blanchard and Perotti, 2002; Mountford and Uhlig, 2009; Ramey, 2011b), the VAR estimates provide a natural benchmark. Moreover, for the medium to long term they tend to be more reliable than estimates based on local projections (Ramey, 2012).

3.1 Model specification

We estimate a panel VAR model on the vector $X_{i,t}$ which includes observations for each quarter, indexed by t, for each country in our sample, indexed by i. In our baseline specification, $X_{i,t}$ includes the log of real government consumption, $g_{i,t}$, the log of real GDP, $y_{i,t}$, and the quarter-on-quarter change of the sovereign default premium, $\Delta s_{i,t}$.[17] The model also includes country-specific constants and time trends, α_i and $\beta_i t$, as well as time-fixed effects, η_t, to control for common factors, such as variations in the price of risk. Importantly, we permit the dynamics of the model to change smoothly, depending on whether the economy operates closer to regime "A" or regime "B". Formally, we rely

[17] We run a number of unit root tests on a country-by-country basis. They suggest the presence of a unit root in case of the default premium, but not in case of real government consumption and real GDP.

on a smooth transition VAR model:

$$X_{i,t} = \alpha_i + \beta_i t + \eta_t + F(z_{i,t})\Lambda_A(L)X_{i,t-1} + [1 - F(z_{i,t})]\,\Lambda_B(L)X_{i,t-1} + \nu_{i,t} \ , \qquad (3.1)$$

where $\Lambda_*(L)$, $* \in \{A, B\}$, is a lag polynomial which features four lags and $\nu_{i,t}$ is a vector of normally distributed errors with regime-dependent covariance matrix $\Omega_t = \Omega_A F(z_{i,t}) + [1 - F(z_{i,t})]\,\Omega_B$.[18]

Auerbach and Gorodnichenko (2012) have popularized smooth transition VAR models for the analysis of the fiscal transmission mechanism. However, while their focus is on how this transmission mechanism changes with the business cycle, we are interested in the impact of "fiscal stress" on fiscal policy transmission. Formally, we weight the polynomials $\Lambda_*(L)$ on the basis of the function $F(z_{i,t})$. We restrict function values to fall in the unit interval and let it be determined by the indicator variable $z_{i,t}$ for which we use the lagged default premium, $s_{i,t-1}$. If fiscal stress is at its maximum, we set $F(z_{i,t}) = 1$, such that the model dynamics depend exclusively on regime "A". Conversely, in the absence of fiscal stress, regime "B" obtains. It is quite unlikely, however, that actual economies operate in either of these two polar regimes. This notion is captured in the estimation, as, for each observation, the impact of the regressors is a weighted average of the dynamics in the two regimes. As a result, all observations contribute to identifying the dynamics that govern in the polar regimes.

The weighting function is critical in this regard. Given that our estimation is based on a fairly large sample, we can rely on the empirical cumulative distribution function of the sovereign default premium in order to determine F, see Figure 2 above. Formally, we have

$$F(z_{i,t}) = \frac{1}{N}\sum_{j=1}^{N} \mathbb{1}_{z_j < z_{i,t}} \ , \qquad (3.2)$$

where, again, $z_{i,t} = s_{i,t-1}$. $\mathbb{1}$ is an indicator function and j indexes all country-time observations, separately for the group of advanced and emerging economies (baseline). Figures A.2 to A.4 in the appendix show how the implied function values vary over time for each country. Alternatively, one may postulate a specific parametric function (e.g.

[18] Using four lags is broadly in line with what information criteria recommend. Results are robust to variations of the lag length.

Auerbach and Gorodnichenko, 2012). Using the empirical CDF as the weighting function has two advantages, however. First, we may remain agnostic and do not impose any parametric restrictions. Second, extreme values are given by observations which actually materialize in sample.[19]

Given the weighting function (3.2) we estimate the VAR model (3.1) by maximizing its likelihood. Because of the nonlinear specification and the large number of parameters to be estimated, we employ the *Tailored Randomized Block Metropolis-Hastings* algorithm to sample from the parameter distribution (Chib and Ramamurthy, 2010).

We compute impulse responses recursively on the basis of the estimated VAR model. In doing so, we kept track of the function value of F, which changes with the endogenous variation of the default premium.[20] Finally, we note that for some variables it is of particular interest to assess their long-run response to fiscal shocks. If a variable is included in $X_{i,t}$ in levels, the impulse response will typically be zero in the long run, because estimating a root to be exactly at 1 is a zero probability event. Instead, by including a variable in first differences, we maintain the hypothesis that its dynamics are governed by a unit root. We may then check whether the cumulative response is significantly different from zero in the long run. Confidence bands are obtained by sampling from the parameter distribution provided by the Metropolis-Hastings algorithm.

3.2 Identification

We follow Blanchard and Perotti (2002) and many others and assume that, within a given quarter, government consumption is predetermined. This assumption is plausible because exhaustive government consumption is unlikely a) to respond automatically to the cycle and b) to be adjusted instantaneously in a discretionary manner by policymakers. To see this, recall that government consumption, unlike transfers, is not composed of cyclical items

[19] One may argue that only governments with relatively large financing needs issue foreign currency bonds and thus appear in our sample. As a consequence, our empirical CDF for fiscal stress may be skewed to extreme observations: those countries with large debt and thus default premium and euro area countries with historically low default premium observations. We check the robustness of our results by also using a parametric logistic transition function and find that they are robust.

[20] We map the default premium into weights according to the CDF for the full sample, because we cannot distinguish between emerging and developed economies in computing impulse responses.

and, in addition, discretionary changes of government spending are subject to decision lags that prevent policymakers from responding to contemporaneous developments in the economy, notably changes in output and the sovereign default premium.

Anecdotal evidence suggests that this holds true also in times of fiscal stress.[21] Still, we cannot rule out that policy measures—while debated for some time—are sometimes spurred by contemporaneous financial-market developments.[22] Against this background, we exploit the specific panel structure of our data set and assess whether policy makers adjust government consumption systematically in response to contemporaneous movements of the sovereign default premium. In particular, in Section 4 we isolate the common component of the sovereign default premium in the cross-section of our panel. We find that it induces the country-specific component of the premium to move strongly on impact, but government consumption with a significant delay only. Hence, we maintain our identification assumption with some confidence.

To impose our identification assumption that government consumption is predetermined relative to the other variables included in the model, we order government spending first in $X_{i,t}$ and equate the first element in $\nu_{i,t}$ with a structural fiscal shock. As a practical matter, we assume a lower-triangular matrix B which maps reduced-form innovations $\nu_{i,t}$ into structural shocks $\varepsilon_{i,t} = B\nu_{i,t}$. We attach no structural interpretation to the other elements in $\varepsilon_{i,t}$.

Influential work by Ramey (2011b) and Leeper et al. (2013) has clarified a potential limitation of our identification strategy: if fiscal policy measures are anticipated by

[21] For instance, in November 2009, European Commission (2009) stated regarding Greece: "in its recommendations of 27 April 2009 . . . the Council [of the European Union] did not consider the measures already announced by that time, to be sufficient to achieve the 2009 deficit target and recommended to the Greek authorities to "strengthen the fiscal adjustment in 2009 through permanent measures, mainly on the expenditure side". In response to these recommendations the Greek government announced, on 25 June 2009, an additional set of fiscal measures to be implemented in 2009 . . . . However, these measures . . . have not been implemented by the Greek authorities so far." In fact, it appears that significant measures were put in place not before 2010Q1, see Greek Ministry of Finance (2010).

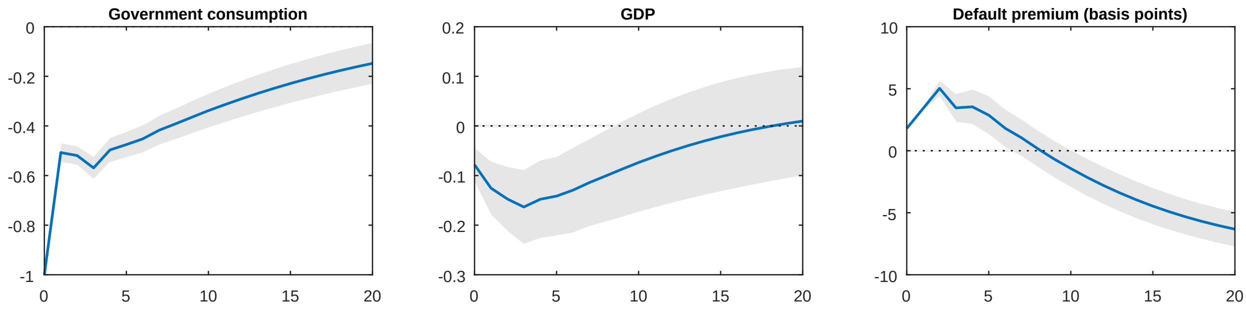
[22] Consider the case of Italy: after some fiscal consolidation in 2010, the default premium kept on rising during the first quarter of 2011 and additional measures were approved by the cabinet on June 30. Finance minister Tremonti, in particular, pushed for severe austerity measures in order to "dispel any spectre of a Greek collapse in Italy" (The Economist, 2011). On July 8 prime minister Berlusconi stated that his finance minister "thinks he's a genius and everyone else is stupid", suggesting some modification to the austerity package. Arguably in response to these remarks, yields on Italian debt rose strongly, such that the package was approved in the Senate without much debate on July 14 (Time Magazine, 2011).

market participants, say, because it takes time to pass legislation or because of other implementation lags, it may fail to uncover the true effect of such measures. Still, because our VAR model includes an inherently forward-looking variable, namely the sovereign default premium, potential problems due to fiscal foresight are likely to be less severe (Sims, 2012). That being said, in Section 4.3, we also consider a variant of our model that features forecast errors of government consumption rather than government consumption itself. As a result, innovations of government consumption are purged of any components which are anticipated by market participants.

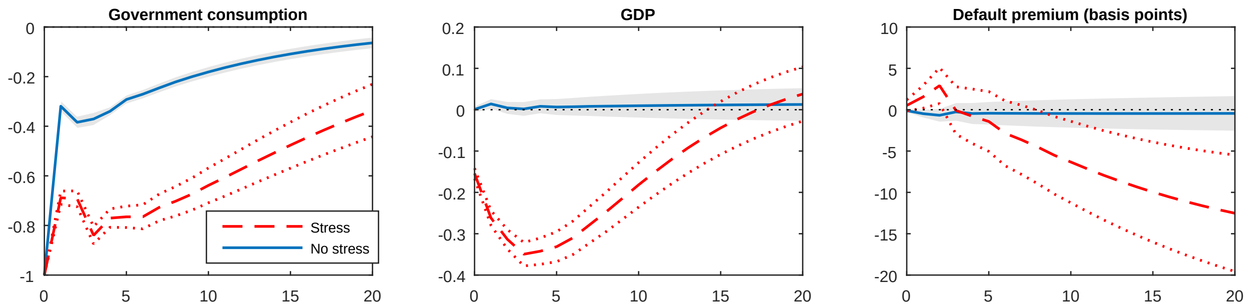
Another popular strategy to identify fiscal shocks is the narrative approach. Following the work of Romer and Romer (2010) for the U.S., Devries et al. (2011) have constructed a data set of fiscal measures for a sample of OECD countries. These fiscal policy measures are identified as being orthogonal to the business cycle on narrative grounds. A large number of these measures are taken in order to reign in public debt or budget deficits. Because sovereign default premia co-move systematically with the latter, such "shocks" are ill-suited to investigate the effect of fiscal policy on the sovereign default premium. Finally, identification based on sign restrictions is not feasible in the context of our analysis, because few of the responses of the variables in our VAR model are uncontroversial as far as fiscal shocks are concerned.

3.3 Results

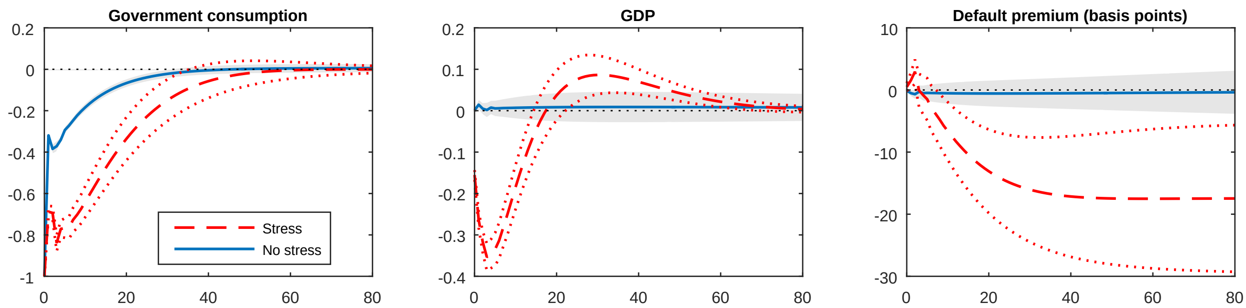
We now report results for our baseline VAR model. Recall that we adopt a parsimonious specification, which features four lags of the endogenous variables real government consumption, real GDP, and the quarter-on-quarter change of the sovereign default premium. Figure 3 displays the impulse response to a cut of government consumption by 1 percent. Here and in what follows, horizontal axes measure quarters, while vertical axes measure the deviation from the pre-shock path, in percent for government spending and output and in terms of basis points for the default premium (for which we display the cumulative response of the quarter-on-quarter change). Solid and dashed lines indicate the point estimates, shaded areas and dotted lines indicate 90 percent confidence bounds. Note



(a) Unconditional VAR (short horizon)



(b) Conditional: fiscal stress vs. no stress (short horizon)



(c) Conditional: fiscal stress vs. no stress (long horizon)

Figure 3: Dynamic response to an exogenous cut of government consumption by 1 percent: unconditional response (top) vs conditional on initial conditions (middle and bottom). Notes: horizonal axis measures quarters, vertical axis measures deviation from pre-shock path in percent and basis points (default premium). Shaded areas and dotted lines indicate 90 percent confidence bounds.

that our baseline estimates are based on group-specific weighting functions (developed vs emerging economies).

Panel (a) shows responses when the VAR model (3.1) is estimated without conditioning on fiscal stress, that is, the coefficient matrices in regimes A and B are restricted to be equal. The response of government spending, shown on the left, is fairly persistent. The maximum output effect obtains in the third quarter after the shock: the decline of output implies a (maximum) multiplier of about 0.7 percent, a value roughly in line with earlier findings (see, for a survey, Ramey, 2011a). The response of the sovereign default premium is shown on the right. The premium increases during the first two years after impact. It declines significantly and below the pre-shock level afterwards. In this sense austerity does pays off—but only in the long run. That a cut of government consumption causes the sovereign default premium to rise in the short run may come as a surprise. Yet we are able to offer a structural interpretation of this finding in Section 5 below. In what follows, we first establish to what extent this result obtains consistently and under which conditions.

We begin with the role of fiscal stress. Specifically, in panel (b) of Figure 3 we show results for the smooth transition VAR model—displaying impulse responses for the limiting cases when there is initially no fiscal stress ($F = 0$, solid lines) and the maximum level of fiscal stress ($F = 1$, dashed lines). The dynamic adjustment differs substantially across the two scenarios. In the absence of fiscal stress, government consumption is less persistent and there is no significant effect on output and the default premium. In contrast, relative to the results shown in panel (a), the output effect is much amplified in case of fiscal stress. The same holds true for the default premium: as in panel (a), it rises initially in the presence of fiscal stress—only to decline significantly below its pre-shock level between one and two years after the shock. Overall, the dynamic adjustment of the economy under fiscal stress resembles that implied by the unconditional estimates. Instead, in the absence of fiscal stress there is not much of an effect.

The results reported in panel (c) are based on the same specification as the results shown in panel (b). Yet in panel (c) we show results for a longer horizon. We observe the default premium plateaus after about 30 quarters and settles on a permanently lower level. Relative to the pre-shock level we find the premium reduced by some 17 basis points in the long-run. This effect is statistically significant, too.

Our VAR model provides an efficient framework to study the long-run effects of fiscal shocks which turn out to conform well with the received wisdom. In what follows we further explore the short-run dynamics, because they are (perhaps) somewhat surprising. For this purpose we rely on local projections, as they offer more flexibility than VARs.

4 Evidence from local projections

We now turn to local projections, as introduced by Jordá (2005). They are quite flexible in accommodating a panel structure and offer a straightforward way to condition the short-run effects of fiscal shocks on the presence of fiscal stress. In addition, local projections allow us to assess to what extent the response of the sovereign default premium to fiscal shocks changes disproportionately in their size. Auerbach and Gorodnichenko (2013), Owyang et al. (2013), and Ramey and Zubairy (2014) also rely on local projections. Their focus, however, is on the fiscal multiplier and on whether it changes with the business cycle and/or the level of nominal interest rates.

4.1 Model specification and identification

In what follows, we are explicit about how the local projection relates to the VAR model employed in Section 3, notably in terms of identification. Specifically, we postulate the same model for government consumption, but do not impose any cross-equation restrictions. We also maintain the identification assumptions introduced in Section 3.2 above. Formally, we rely on the following model:

$$g_{i,t} = F\left(z_{i,t}\right)\Gamma_A(L)X_{i,t-1} + \left[1 - F\left(z_{i,t}\right)\right]\Gamma_B(L)X_{i,t-1} + \varepsilon^g_{i,t} \ . \tag{4.1}$$

Here $X_{i,t}$ is a vector of regressors with $g_{i,t} \in X_{i,t}$ and, as before, $\Gamma_A(L)$ and $\Gamma_B(L)$ are lag polynomials of coefficient matrices capturing the dynamic effect of the regressors in each regime. Deterministic terms are omitted to simplify the exposition. Under the assumption that government consumption is predetermined, $\varepsilon^g_{i,t}$ is a structural innovation to government consumption.

Eventually, we are interested in the dynamic effects of this innovation. Formally, letting $x_{i,t+h}$ denote the response of a particular variable at horizon h to an innovation at time t, we seek to retrieve the following relation:

$$x_{i,t+h} = \alpha_{i,h} + \beta_{i,h}t + \eta_{t,h} + F\left(z_{i,t}\right)\psi_{A,h}\varepsilon^g_{i,t} + \left[1 - F\left(z_{i,t}\right)\right]\psi_{B,h}\varepsilon^g_{i,t} + u_{i,t+h} \ . \tag{4.2}$$

Here $\alpha_{i,h}$ and $\beta_{i,h}t$ are a country-specific constant and a country-specific trend, respectively, and $\eta_{t,h}$ captures time-fixed effects. Also, at each horizon, the response of the dependent variable to the fiscal innovation is allowed to differ across regimes "A" and "B", with the ψ-coefficients on the $\varepsilon_{i,t}^{g}$ terms indexed accordingly. $F\left(z_{i,t}\right)$ captures fiscal stress as explained in Section 3.1. The error term $u_{i,t+h}$ is assumed to have a zero mean and strictly positive variance.

The innovation $\varepsilon_{i,t}^{g}$ is generally not observable. We therefore rearrange (4.1) and substitute in (4.2) to obtain the following regression equation

$$
\begin{aligned}
x_{i,t+h} = \alpha_{i,h} + \beta_{i,h}t + \eta_{t,h} & \\
& + F\left(z_{i,t}\right)\psi_{A,h}g_{i,t} + \left[1 - F\left(z_{i,t}\right)\right]\psi_{B,h}g_{i,t} \\
& - \psi_{A,h}\Gamma_{A,h}(L)\left[F\left(z_{i,t}\right)\right]^2 X_{t-1} - \psi_{B,h}\Gamma_{B,h}(L)\left[1 - F\left(z_{i,t}\right)\right]^2 X_{t-1} \\
& - \psi_{A,h}\Gamma_{B,h}(L)F\left(z_{i,t}\right)\left[1 - F\left(z_{i,t}\right)\right]X_{t-1} - \psi_{B,h}\Gamma_{A,h}(L)F\left(z_{i,t}\right)\left[1 - F\left(z_{i,t}\right)\right]X_{t-1} \\
& + u_{i,t+h}\ .
\end{aligned}
\tag{4.3}
$$

We estimate (4.3) using OLS where, in order to improve the efficiency of the estimates, we include the residual of the local projection at $t+h-1$ as an additional regressor in the regression for $t+h$ (see Jordá, 2005). For each forecast horizon, the sample is adjusted accordingly to use all available country-quarter observations.

Projection (4.3) directly captures the dynamic effects of a fiscal innovation conditional on the circumstances under which it occurs. Formally, the response in period $t+h$ to a government consumption impulse in period t, $\varepsilon_{i,t}^{g}$, conditional on the economy experiencing a particular state today, indexed by $z_{i,t}$, is given by the regression coefficients on $g_{i,t}$ in equation (4.3):

$$
\left.\frac{\partial x_{i,t+h}}{\partial g_{i,t}}\right|_{z_{i,t}} = F\left(z_{i,t}\right)\psi_{A,h} + \left[1 - F\left(z_{i,t}\right)\right]\psi_{B,h}\ . \tag{4.4}
$$

This expression illustrates that computing impulse responses based on a single-equation approach does not require us to make additional assumptions on the economy staying in a particular regime (see also the discussion in Ramey and Zubairy, 2014). Rather, the local projection at time t directly provides us with the average response of an economy in state $z_{i,t}$ going forward. Note also that equation (4.4) is just a linear combination of regression coefficients. We can thus rely on a Wald-type test to assess whether responses at a particular horizon are significantly different from each other as a result of different initial conditions.

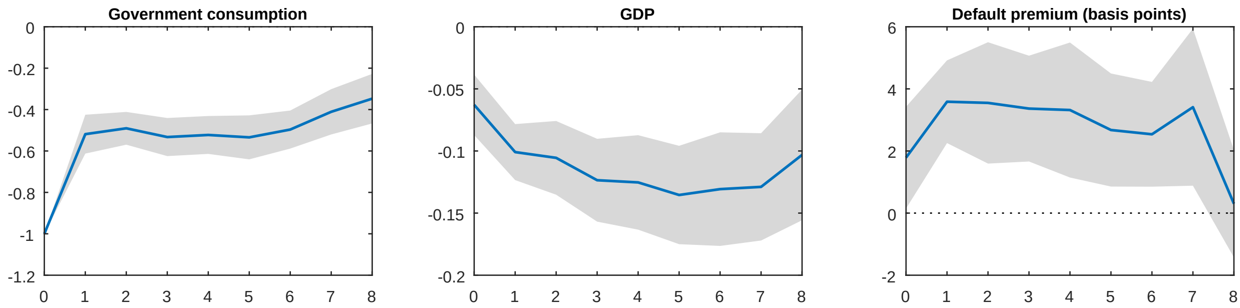
4.2 Projections on government consumption

We estimate the local projection (4.3) on the full sample, using again four lags of government consumption, GDP, and the default premium.[23] We obtain the impulse responses to a shock of government consumption for all three variables, allowing for a maximum horizon of $h = 8$ quarters. Extending the horizon comes at the expense of degrees of freedom in the time-series dimension and estimates tend to be less reliable at longer horizons (Ramey, 2012). The size of the shock corresponds to one percent of government consumption. Figure 4 shows the results. As before, shaded areas and dotted lines indicate 90 percent confidence bounds, based on standard errors which are robust with respect to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

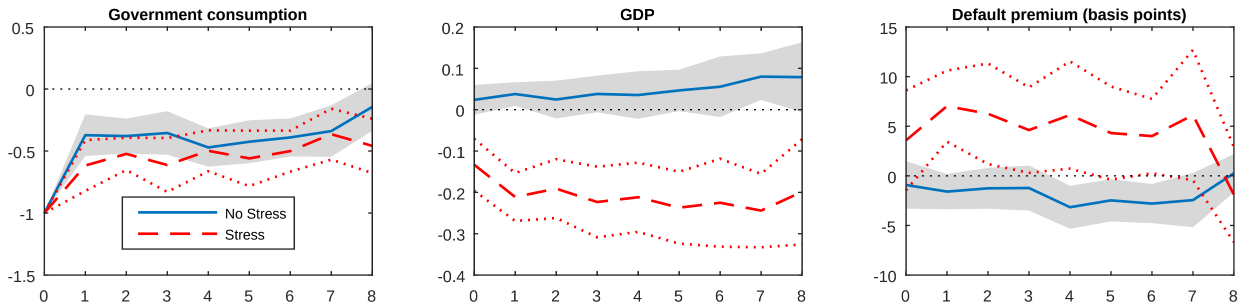
Panel (a) of Figure 4 displays the estimates obtained without conditioning on fiscal stress. Government consumption, shown on the left, remains depressed for an extended period, but eventually returns to its pre-shock level. The response of GDP, displayed in the middle, declines initially by about 0.1 percent, and more strongly thereafter. The strongest effect of about -0.15 obtains after roughly 1.5 years. On the right, we show the response of the default premium: it increases in response to the cut of government consumption. The impact and maximum response is about 2 and 4 basis points, respectively. Overall, the results are quite similar to the unconditional effects obtained from the VAR model, displayed in panel (a) of Figure 3.

Panel (b) shows results conditional on initial conditions in terms of fiscal stress. Solid lines represent point estimates, provided that there is no fiscal stress. Dashed lines represent the results conditional on the presence of fiscal stress. It turns out that, as

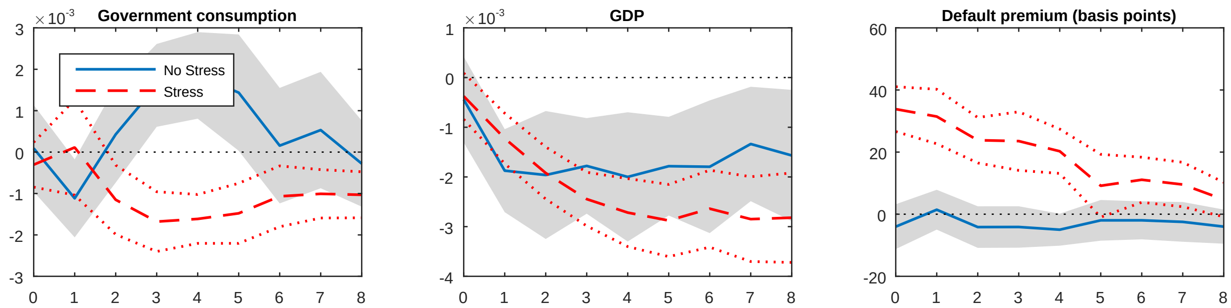
[23]We include the default premium in levels, as we focus on the short run. Using first differences instead yields almost identical results for the implied response of the default premium.



(a) Response to cut of government consumption: unconditional



(b) Response to cut of government consumption: conditional on initial conditions



(c) Response to increase of common factor of default premium: conditional on initial conditions

Figure 4: Dynamic responses. Notes: panels (a) and (b) show impulse responses to cut of government consumption by 1 percent; panel (c) shows responses to 1 percentage point increase of the common factor of the sovereign default premium. Horizonal axis measures quarters, vertical axis measures deviation from pre-shock path in percent and basis points (default premium). Shaded areas and dotted lines indicate 90 percent confidence bounds.

with the VAR estimates, differences across regimes are rather stark. Neither output nor the default premium respond much to the shock in the absence of fiscal stress. Output, however, falls strongly and the premium rises sharply if government consumption is cut in times of fiscal stress. The impact and peak responses of the default premium are summarized in Table 3. We also check whether the responses in both regimes are statistically different from each other. According to a Wald test, the null hypothesis of an equal response can generally be rejected for output and the default premium at all

Table 3: Impact, peak, and long-run response of default premium (basis points)

	Local projection		VAR	
	Impact	Peak	Short-run peak	Long run
Unconditional	2	4	5	-6
Fiscal stress	4	7	3	-17
No stress	-1	-3	-1	1

Notes: Response to a 1 percent cut of government consumption. The short-run peak is the maximum of the absolute value of the response of the default premium during the first 8 quarters.

horizons. The null is not rejected for government consumption.

Table 3 also shows that our main result is robust across different model specifications: a cut of government consumption raises the sovereign default premium in the short run, provided that fiscal stress is high. To the extent that this finding is surprising, it may raise doubts about the identification assumption which we entertain throughout, namely, that government consumption is predetermined within the quarter. After all, results might be driven by reverse causality: as the sovereign default premium rises, governments may immediately cut public consumption to calm financial markets.

The panel structure of our data set allows us to assess this conjecture formally. For this purpose, we first extract a common factor in the default premium along the cross-sectional dimension of our panel by means of a principal component analysis (see Longstaff et al., 2011, for a similar approach). We conduct the analysis separately for developed and emerging economies.[24] In a second step, we run a local projection on our panel, relating current and future government consumption in each country (as well as output and default premium) on the common factor. Here identification rests on the assumption that the common factor—variations of which may, for instance, reflect changes in the stochastic discount factor of global investors—is not contemporaneously affected by country-specific developments.

We condition the effects on the presence of fiscal stress as above and show results for

[24]As a practical matter, to deal with the missing values in our panel, we conduct a "Probabilistic principal component analysis" as implemented in Matlab 2015b's `ppca`-command. To find the global mode of the likelihood function, we run the analysis with random starting values and keep those for which we obtain the highest likelihood. The sample starts in 1994Q4 for emerging economies and in 1991Q4 for developed economies. We exclude time-fixed effects from the local projections including the common factor.

an increase of the common factor by 1 percentage point in panel (c) of Figure 4. There is no significant response of government consumption in the very short run. If there is fiscal stress, government consumption is reduced, as the common factor increases, but only with a delay of 2 or 3 quarters. We explore this finding in more detail in ongoing work (Born et al., 2016). The delayed response of government consumption is particularly remarkable, because output and the sovereign default premium tend to respond instantaneously to the common factor. In light of these results, we judge concerns that our main finding—the short-run response of the default premium to a cut of government consumption—is the result of reverse causality as unwarranted.

4.3 Projections on forecast errors of government consumption

Our baseline model assumes that innovations of government consumption are a surprise to market participants. To the extent that this assumption fails to be satisfied in actual time-series data, estimates are biased (Leeper et al., 2013; Ramey, 2011b). To address this issue we follow Ramey (2011b) and construct a measure of fiscal shocks which is purged of possible anticipation effects, namely the forecast error of government consumption.

The OECD compiles semiannual forecasts of government consumption, covering the period from 1986 to 2014 for an unbalanced panel of OECD countries. Forecasts are prepared at the end of an observation period, namely, in June and December of each year and tend to perform quite well (Auerbach and Gorodnichenko, 2012). In what follows, we aggregate the quarterly observations in our sample to obtain observations at semi-annual frequency. We evaluate the weighting function F using the end-of-period values of the (lagged) default premium. Regarding the forecast errors of government consumption, we compute growth rates rather than levels, because the OECD changes the base year several times during our sample period. Our transformed sample consists of 697 semi-annual observations.

We estimate the local projection (4.3) on this sample, replacing the level of government consumption $g_{i,t}$ with the period-t forecast error of the growth rate of government consumption, while still including the lags of government consumption as controls. The

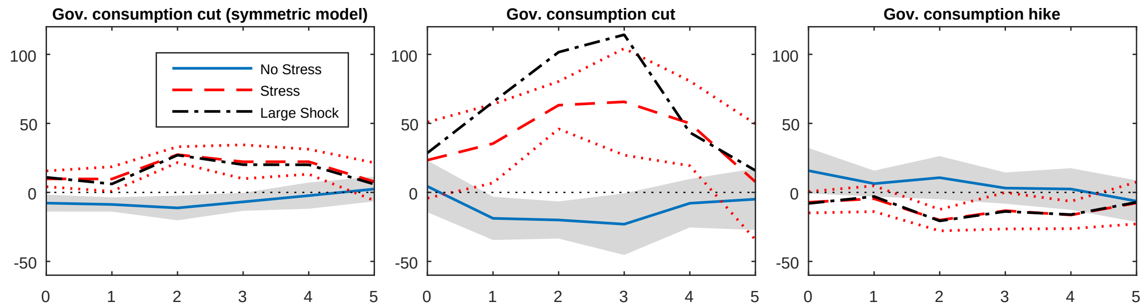


Figure 5: Dynamic response of sovereign default premium to change of government consumption: projection on forecast errors using semi-annual observations. Horizontal axis measures half years. Solid lines: responses conditional on absence of fiscal stress; dashed lines: fiscal stress; dash-dotted lines: response conditional on change of government consumption larger than 2 standard errors (assuming fiscal stress). Shock size normalized to one percent of government consumption in all cases. Left: symmetric model; middle: government consumption decrease; right: government consumption increase.

response of the sovereign default premium is shown on the left of Figure 5. It is quite similar to that shown in Figure 4, despite of differences in the sample, the sampling frequency, and despite of the correction for anticipation effects.[25]

Given our findings for fiscal stress, one may wonder whether raising government consumption is a means to reduce the sovereign default premium. After all, local projection (4.3) restricts the responses to increases and decreases of government consumption to be symmetric. In what follows, we relax this restriction and estimate (4.3) while treating spending cuts and spending hikes as distinct regressors. Figure 5 shows the results: the response to a spending cut is shown in the middle, while the response to a spending increase is shown on the right. Allowing for asymmetric effects turns out to be important: in times of fiscal stress, the default premium responds strongly only if government consumption is cut. The response to a spending increase, on the other hand, is muted—both in the presence of fiscal stress and in the absence thereof.

The effects of changes of government consumption may more generally depend on its size (Giavazzi et al., 2000). In order to assess this possibility we include an interaction term in the local projection (4.3): it interacts the shock with a dummy variable which is

[25]Figure A.5 in the appendix contrasts the results for forecast errors and for government spending based on an identical sample. Results are very similar across both specifications, also from a quantitative point of view. This confirms earlier findings by Beetsma and Giuliodori (2011), Born et al. (2013), and Corsetti et al. (2012b).

Table 4: Response of sovereign default premium (basis points), forecast error identification, semi-annual observations

	Baseline: spending cut in symmetric model		
	Stress	No stress	Difference
Impact	9	-8	17*
Maximum (stress)	31	-14	46*

	Controlling for sign: spending cut		
	Stress	No stress	Difference
Impact	23	4	19
Maximum (stress)	66	-23	89*

	Controlling for sign: cut vs hike (fiscal stress)		
	Spending cut	Spending hike*(-1)	Difference
Impact	23	7	16
Maximum (cut)	66	13	46*

	Controlling for size: baseline vs large spending (fiscal stress)		
	Spending cut	Large spending cut	Difference
Impact	23	29	6
Maximum (large cut)	66	114	48*

Notes: Results based on model (4.3) and on sample for which forecasts of government consumption are available; shock size normalized to one percent of government consumption in all cases (see footnote (26)). Maximum response under stress/cut/large cut (> 2 SE) is compared to response in the absence of stress/increase/cut at same horizon. An asterisk indicates significance at the 5 percent level. In the bottom panel, the test of significance of the difference is based on the model which includes an interaction term (see main text).

one whenever the shock is larger than two standard deviations (on a country-by-country basis). We focus our analysis on the case where fiscal stress is present and find that size indeed matters. Large spending cuts raise the default premium more than proportionally, as illustrated by the dashed-dotted line in panel (b) of Figure 5.[26] The figure also shows that the effect is not present in the symmetric model or in case of spending hikes.

Table 4 summarizes how changes of government consumption impact the default premium. Importantly, it shows the difference of the response due to various conditioning

[26]To ensure comparability across different shock sizes, all panels report IRFs to a shock of 1 percent of government consumption, although the average shock size is actually different.

factors, both for the impact and the maximum effect. First, the premium rises in the short run only if there is “fiscal stress” (upper panel). Second, while cutting government consumption in times of fiscal stress increases the default premium, a spending hike in this situation leaves it basically unchanged (second panel). Third, under fiscal stress, large austerity programs have a more than proportional effect on the premium in the short run (fourth panel). These differences are statistically significant in most cases (always when considering the maximum effects).

4.4 Sensitivity analysis

We explore the robustness of our findings across a range of alternative specifications and subsets of country-time observations. In what follows we briefly discuss results and relegate figures to the appendix. Unless stated otherwise, our point of departure is the local projection (4.3) with government consumption as regressor, since the sample is largest for this specification.

Additional variables

Our baseline specification is parsimonious, because earlier studies of the fiscal transmission mechanism suggest that including additional variables in the model has little effect on the results (e.g., Born and Müller, 2012). In what follows we nevertheless include a number of additional variables in the local projection. In doing so, we not only assess the robustness of our results, but also shed further light on the transmission mechanism. A first set of additional variables includes the debt-to-GDP ratio, government net lending relative to GDP, private consumption, and private investment. We include each variable, in turn, in $X_{i,t}$ and project them on government consumption. Note that due to data availability, the number of observations is somewhat reduced relative to the baseline.

The panels in Figure A.6 in the appendix show results. On the left and in the middle of each panel, we show the impulse responses of output and the sovereign default premium, respectively. They turn out to be quite similar to those obtained for the baseline model. On the right of each panel, we show the impulse responses of the additional variable.

Responses vary considerably depending on whether there is fiscal stress or not. The responses of debt and the government budget are particularly informative: debt relative to output rises in response to a spending cut, if fiscal stress is high. It declines, albeit very gradually, in the absence of stress. Similarly, after a cut of government consumption, net lending increases relative to output only in the absence of fiscal stress.

We also consider variations of the baseline model which include, in turn, confidence and stock market returns as additional variables. Confidence data is provided by the Ifo World Economic Survey (WES), which surveys a number of experts for all countries in our sample.[27] Earlier research on the consequences of fiscal consolidations has argued that its impact on "confidence" is crucial (see, for instance, the discussion in Perotti, 2013). Bachmann and Sims (2012) find that confidence responds strongly to fiscal shocks during periods of economic slack. Also, both confidence and stock market returns are forward-looking variables and may thus control for fiscal foresight. Results are shown in Figure A.7: the responses of output and the default premium are again fairly unchanged relative to the baseline model. The responses of confidence and stock market returns are not significantly different from zero—irrespective of initial conditions.

Cross-sectional heterogeneity

The central theme of our analysis is that the effects of fiscal policy differ along a number of important dimensions. Still, even after accounting for these (as well as for country and time-fixed effects) there may still be important cross-sectional heterogeneity across countries which is unaccounted for in our baseline model. As a first step to address this concern, we consider a number of sample splits: a sample that includes only euro area countries, a sample of euro area periphery countries which were hit hardest by the crisis (Greece, Ireland, Italy, Portugal, Slovenia, Spain), and a sample of the remaining euro area countries.

Results, shown in Figure A.8, tend to be qualitatively similar to those obtained for the

[27] Respondents are asked to classify their expectations for the next six months using a grid ranging from 1 (deterioration) to 9 (improvement). 5 indicates that expectations are "satisfactory" (see, e.g., Kudymowa et al., 2014).

full sample—notably in terms of the differential impact of fiscal stress. The same holds for sub-samples comprising developed and emerging economies only, see panels (a) and (b) of Figure A.9. As a caveat, however, we note that there are sizeable differences in some instances, partially reflecting a strong decline in sample size. Similarly, we rule out that results are driven by the Great Recession and report estimates for the pre-2007Q3 period in panel (c) of Figure A.9.

A second, more formal approach to control for cross-sectional heterogeneity is to employ a mean-group estimator (Pesaran and Smith, 1995). In this case we estimate model (4.3) for each country separately and average over the cross section of coefficients afterwards. As this leaves us with few degrees of freedom, we only estimate the model without conditioning on fiscal stress and use the panel model with slope homogeneity as our benchmark. Figure A.10(a) shows that our results are robust to allowing for cross-sectional heterogeneity in parameters.

Independent monetary policy

Some observers have argued that the sovereign default premium, notably during the recent euro area crisis, is driven by "market sentiment" rather than "fundamentals" . According to a popular narrative, the fact that euro area countries have surrendered monetary independence is crucial in this regard (see, e.g., De Grauwe and Ji, 2012). Independent central banks, so the argument goes, can act as a lender of last resort to governments and thereby rule out speculative runs on governments (see Farhi et al., 2013, for a formal treatment). Hence, whether a central bank is independent or not may matter for the dynamics of the default premium, at least the one paid on domestic-currency debt. Yet, by reducing the likelihood of runs on domestic debt, it is likely that there are spillover effects on the default premium paid on foreign-currency debt, too.

To explore whether this aspect matters for our results, we identify those countries in our sample which are either members of a monetary union or have officially dollarized.[28] Figure A.10 shows the results for this group of countries in panel (b) and results for

[28] Ecuador since 2000Q1 and El Salvador since 2001Q1 use the dollar as their official legal tender (see Levy Yeyati and Sturzenegger, 2002).

countries with their own legal tender in panel (c). We find that conditioning on monetary independence has little bearing on our results, although the response of the default premium tends to be weaker in countries with their own legal tender.

Sample without IMF program countries

Austerity programs are frequently part of the conditionality of IMF assistance which is typically called upon when the sovereign default premium is high. To ensure that our results are not driven by such episodes, we drop all observations for which a country qualifies as an IMF "program country". We rely on the IMF's "History of Lending Arrangements" and classify countries as "program countries" if there is either a "Standby Arrangement" or an "Extended Fund Facility". Results are quite similar to those for the baseline sample. They are shown in panel (d) of Figure A.10.

Measurement of default premium

In our baseline specification we measure the default premium in basis points. Benign times are effectively characterized by a premium of close to zero. Impulse responses computed for the absence of fiscal stress may therefore imply that the premium becomes potentially negative. Economically this makes little sense. We therefore consider an alternative specification where the premium is measured in logs. The results, shown in Figure A.11(a), are qualitatively similar, with the premium in the absence of fiscal stress roughly constant.

Data quality

Our sample includes observations for European countries in the 1990s. These data do not fully meet the more recent standards for the compilation of quarterly non-financial accounts of the government (see, e.g., Eurostat, 2011). As a robustness check, we therefore estimate our model on a conservative subsample where the data quality is higher.[29] Results

[29] We checked with national statistical agencies and adjusted the Ilzetzki et al. (2013) sample where necessary. Using this conservative sample eliminates about 10% of our observations for developed economies.

are shown in Figure A.11(b). They are very similar to our baseline sample.

Boom and recessions

Times of fiscal stress are most likely times of low output growth. The converse is not true: a recession does not necessarily give rise to fiscal stress. Still, to put our results into perspective, it is useful to assess to what extent the effects of fiscal shocks on the default premium change with the state of the business cycle. For this purpose, we estimate a variant of model (4.3) while conditioning on the state of the cycle (rather than fiscal stress). In this case we follow Auerbach and Gorodnichenko (2013) and use a measure of the output gap as indicator variable z_{it}. However, in contrast to their analysis, we rely on the empirical cumulative distribution function to weigh regressors.[30]

Figure A.11(c) shows the results. In line with earlier findings by Auerbach and Gorodnichenko (2013), we find that the output effects of fiscal policy are considerably stronger during recessions. We also find that the default premium increases during recessions and falls during booms in response to cuts of government consumption. We thus obtain a pattern of responses quite comparable to those once we condition on fiscal stress. Perhaps surprisingly, while conditioning on fiscal stress and recessions yields very similar results, we find that the overlap of stress and recession episodes is far from complete. In particular, the correlation of the empirical CDF is only moderate (see Table A.1 and Figures A.2 to A.4 in the appendix.).

5 Interpretation

Our empirical analysis reveals a robust pattern: a cut of government consumption raises the sovereign default premium in the short run—provided that fiscal stress is high. We rationalize this finding within a structural model of sovereign default à la Arellano (2008). In order to keep the analysis as transparent and focused, we modify the original model as

[30] To obtain a measure for the output gap, we compute a five-quarter moving average of the first difference of log output. The resulting series is then z-scored and filtered using an Hodrick-Prescott filter with smoothing parameter $\lambda = 160{,}000$. This is the value used in Auerbach and Gorodnichenko (2013), adjusted for our quarterly sample following Ravn and Uhlig (2002).

little as possible. Specifically, to capture key aspects of our empirical setup, we depart from the original model by allowing for (a) exogenous variations in government spending and (b) a multiplier effect of such variations on output. The model features stationary dynamics, as all variables converge to an ergodic mean. It is thus suited for an analysis of the short-run dynamics only. In what follows, we briefly outline the small open economy.

The government engages in intertemporal trade, as it seeks to maximize the expected utility of the representative household given by

$$E_0 \sum_{t=0}^{\infty} \beta^t U(c_t) \ . \tag{5.1}$$

Here, $0 < \beta < 1$ is the discount factor and c_t is private consumption. Output y_t is given by

$$y_t = \bar{y} e^{\mu \hat{g}_t} \ , \tag{5.2}$$

where $\bar{y}$ is a positive constant and $\hat{g}_t$ is the percentage deviation of government consumption from its long-run value $\bar{g}$. We abstract from variations in output due to other factors. Government consumption varies exogenously and impacts household utility additively separable from c_t. Hence, we omit it from the exposition in (5.1). Parameter μ in the expression above is given by $(\bar{g}/\bar{y})\epsilon$, where ϵ measures the fiscal multiplier, that is, the percentage change of output, given an increase of government consumption equal to one percent of GDP.[31]

The government sells debt to risk-neutral, international investors. It cannot commit to repay, but decides in a discretionary manner on whether to service the outstanding debt or not in each period. In case it repays, the flow budget constraint of the economy is given by

$$y_t + (1 + r_t)^{-1} d_{t+1} - d_t = c_t + \bar{g} e^{\hat{g}_t} \ , \tag{5.3}$$

where d_t is beginning-of-period debt, d_{t+1} is newly issued debt which pays one unit of output in the next period if it is redeemed. It trades at a discount $(1 + r_t)^{-1}$. The absence

[31] We skip a possible microfoundation of the multiplier based on, for instance, endogenous labor supply or a working capital constraint in order keep the analysis tractable. Mendoza and Yue (2012) develop a model of optimal sovereign default and endogenous output determination.

of arbitrage possibilities requires the following condition to be satisfied

$$1 + r_t = \frac{1 + r}{1 - \delta_t} \ . \tag{5.4}$$

Here δ_t is the probability of default and r is a risk-free return which international investors earn elsewhere. Investors are assumed to form expectations rationally, that is, they understand the determinants of the default decision which we discuss below. We also assume that investors do not receive any payment in case of default. The sovereign default premium is given by $r_t - r$.

In the event of default the country is excluded from international financial markets. It may be allowed to reenter financial markets with probability θ in each period thereafter. In addition, there is an asymmetric output cost, such that output is given by

$$y_t^{\text{def}} = \min(y_t, \bar{y}^{\text{def}}) \tag{5.5}$$

as long as the country remains in the default state. Here $\bar{y}^{\text{def}}$ is a constant defining the maximum output level. Consumption in the default state is then given by

$$c_t^{\text{def}} = y_t^{\text{def}} - \bar{g}e^{\hat{g}_t} \ . \tag{5.6}$$

To characterize the decision problem of a government that enters the current period with debt d_t and government spending g_t, it is useful to define the value of having the option to default, $v^o(d_t, g_t)$, as follows

$$v^o(d_t, g_t) = \max_{\{c_t, \text{def}\}} \left\{ v^c(d_t, g_t), v^{\text{def}}(g_t) \right\} \ . \tag{5.7}$$

Here, $v^c(d_t, g_t)$ is the continuation value associated with not defaulting, while $v^{\text{def}}(g_t)$ is the value of repudiating debt. Setting $d_{t+1} = 0$, is is defined recursively as

$$v^{\text{def}}(g_t) = U(c_t^{\text{def}}) + \beta \int_{g_{t+1}} \left[\theta v^o(0, g_{t+1}) + (1 - \theta) v^{\text{def}}(g_{t+1}) \right] f(g_{t+1}, g_t) dg_{t+1} \ . \tag{5.8}$$

Here g_{t+1} denotes next period's government consumption. The continuation value of not defaulting, in turn, is given by

$$v^c(d_t, g_t) = \max_{\{d_{t+1}\}} \Bigg\{ u\left(y_t + q(d_{t+1}, g_t)d_{t+1} - d_t - g_t\right) \\ + \beta \int_{g_{t+1}} v^o(d_{t+1}, g_{t+1}) f(g_{t+1}, g_t) dg_{t+1} \Bigg\} . \tag{5.9}$$

Hence, exactly as in Arellano (2008), the government decides on the optimal level of borrowing and on whether to repay in order to maximize household utility. In doing so, it is also constrained not to run Ponzi schemes. We skip the definition of a recursive equilibrium, because it is isomorphic to the one in Arellano (2008).

Also, rather than providing a full-fledged analysis of the model, we focus on how the default premium reacts to exogenous variations of government consumption. For our model simulation, we assume $U(c_t) = c_t^{1-\sigma}/(1-\sigma)$ as in Arellano (2008) and postulate an AR(1)-process for government consumption:

$$\hat{g}_t = \rho_g \hat{g}_{t-1} + \varepsilon_t^g,\ \varepsilon_t^g \sim \mathcal{N}(0, \sigma_g^2) . \tag{5.10}$$

In assigning parameter values, we assume that a period in the model corresponds to one quarter and adjust the parameter values of Arellano's calibration to annual frequency accordingly. In terms of government consumption, we assume $\bar{g} = 0.18$, the average share of government consumption in our sample. We set $\rho_g = 0.986$ to capture the persistence of a shock to government consumption under fiscal stress and $\sigma_g = 0.014$ to match the standard deviation of a government consumption shock in our sample. For the government spending multiplier we assume $\epsilon = 0.7$, a value, as discussed above, in line with the estimates reported in the literature. It is also consistent with what our empirical analysis suggests for times of fiscal stress. Table 5 summarizes the parameter values. The model is solved by discretizing the AR(1)-process into a 101-point Markov chain in the range of $\pm 4\ \sigma_g$ and using value function iteration on a 2000-point grid for debt on $[0, 1.32]$.

We compute the generalized impulse response of the default premium to a change in government consumption.[32] As in our empirical analysis we distinguish initial conditions

[32] Generalized impulse responses account for the nonlinearity of the model (see, Koop et al., 1996). We compute responses on the basis of stochastic simulations as the difference in the dynamics after a deterministic shock at time 1 and the dynamics in the absence of a shock. We approximate the sovereign default premium $r_t - r$ by δ_t. As in our empirical analysis we exclude default episodes. We report results

Table 5: Parameter values used in model simulations

Parameter	r	σ	β	θ	$\bar{y}$	$\bar{y}^{\text{def}}$	$\bar{g}$	ρ^g	σ_g	ϵ
Value	1.7%	2	0.988	0.07	1	0.969	0.18	0.986	0.014	0.7

Note: Parameter values follow Arellano (2008) where applicable (adjusted for model calibration to quarterly frequency). See text for details.

in terms of fiscal stress. For this purpose we adjust the initial level of debt. For a debt level of 8% below the average, fiscal stress is absent, that is, the default premium is zero. To capture fiscal stress, we assume that debt is initially 8 percent above the average level. In this case the default premium amounts to 36 basis points.[33]

Results are shown in Figure 6: solid lines represent the responses in the absence of fiscal stress, while dashed lines correspond to the case of fiscal stress. Horizontal axes measure time in quarters, vertical axes measure basis points. On the left and in the middle, we show results for an increase and a decrease of government consumption by one standard deviation, respectively. On the right, we consider a decrease by three standard deviations. To make responses comparable, we divide the response by 3 in case of the large shock.

The premium is hardly affected by the shock in the absence of stress, even if the shock is large. Instead, it reacts strongly if there is fiscal stress. The response is still relatively weak, however, if government consumption increases (left). Instead, the premium rises strongly, if government consumption is reduced (middle). Moreover, the response of the premium increases more than proportionally in the size of the shock (right). In sum, the model predictions are in line with our empirical findings. This holds true for the role of initial conditions in terms of fiscal stress, as well as for the role of the sign and the size of an innovation of government consumption. Moreover, it also holds true for the finding that the premium rises in response to a cut of government consumption.

The predictions of the model depend on the assumption that there is a positive multiplier effect.[34] To see why, recall that a cut of government consumption reduces

based on the average over 100,000 replications.

[33]For higher debt levels, in case of a large (3 standard deviation) cut of government consumption default turns out to be optimal.

[34]If we set $\epsilon = 0$, a reduction of government consumption reduces the sovereign default premium.

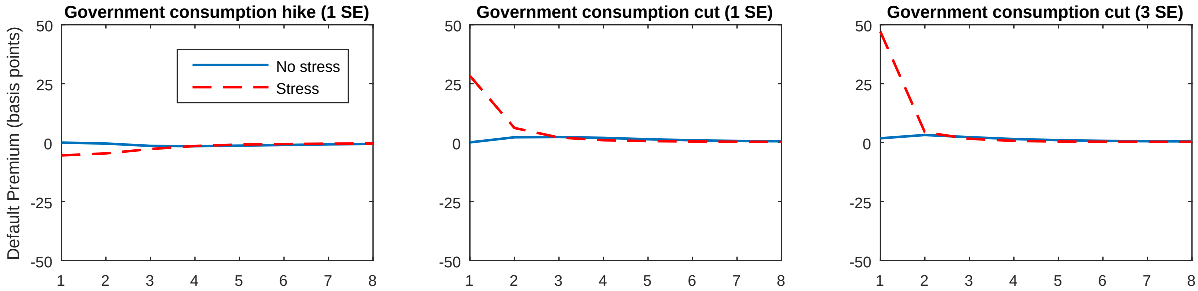


Figure 6: Generalized impulse response to a change of government consumption. Notes: Horizontal axes represent quarters. Vertical axes represent deviations from the path w/o shock, measured in basis points. Right: response is divided by 3.

output due to the multiplier effect. Yet, because the multiplier is below unity, output net of government consumption increases in response to a cut of government consumption. This, all else equal, makes default less attractive. The effect is muted, however, relative to the scenario without a multiplier effect. Moreover, there is a second margin by which the multiplier effect impacts the default decisions: as output declines with government consumption, the output loss due to default declines as well. This is because, in specifying the kinked function (5.5), we follow Arellano and assume that default is less harmful if the economy is already in dire straits.

Overall, the cut of government consumption raises the incentive to default. This becomes apparent from the increase of the default premium which reflects the assessment of market participants. They understand the trade-off which determines the default decision and price default risk in an actuarially fair manner, namely according to condition (5.4). Fiscal stress amplifies the response of the premium, because condition (5.4) is convex in δ_t. For the same reason, the premium increases disproportionately as government consumption is reduced. We thus stress that sovereign debt is priced consistently via condition (5.4). Put differently, there is nothing "schizophrenic" about financial markets: in the presence of fiscal stress, a cut of government consumption may trigger a further increase of the default premium, because markets correctly foresee a heightened temptation to default. This temptation lasts as long as output remains depressed.

6 Conclusion

Does austerity cause the sovereign default premium to decline? In pursuing this question, we make two distinct contributions. First, we set up a new data set containing data for 38 emerging and advanced economies. We assemble quarterly observations for an unbalanced panel from 1990 to 2014, not only for the sovereign default premium, but also for government consumption and output. A first look at the data allows us to establish a number of basic facts. First, while there is a large variation in the default premium, both across time and countries, it is moderate for the largest part of our sample. Second, the default premium is strongly countercyclical. The correlation of the default premium and output growth is negative in all 38 countries. Third, across countries there is no systematic correlation pattern for the default premium and government consumption.

As a second contribution, we assess how the default premium responds to a cut of government consumption. In doing so, we account for three dimensions. First, we distinguish the short and the long run. We find that a cut of government consumption raises the premium in the short run, but reduces it in the long run. Second, we condition on initial conditions. We find that the short-run increase of the premium obtains only in the presence of fiscal stress, that is, if the default premium is already high. Third, the effect of an increase of government consumption depends on its size: the premium reacts more than proportionally, as the cut of government consumption gets larger.

We rationalize our findings for the short run within a structural model of optimal sovereign default. It turns out that assuming a fairly moderate fiscal multiplier is sufficient for the model predictions to conform with the evidence. In the model, risk-neutral investors price sovereign default risk in an actuarially fair manner. They demand a higher default premium in response to a cut of government consumption, because they understand that the temptation to default increases as output is temporarily reduced. This is also consistent with our findings for the long run: the premium peaks after about one year and comes down rather quickly afterwards—at about the same time as output rebounds.

Our results have important implications for policy. First, whether austerity is painful in the short run depends on the level of fiscal stress. As spending cuts cause little harm under benign initial conditions, it is advisable to prevent fiscal stress from building up in the first place. Admittedly, more often than not policy makers will have missed this opportunity. As a result, austerity is likely to be painful in the short run and the temptation to renege on debt obligations increases. Because market participants understand this, they demand a higher default premium. A naive observer may therefore conclude that "austerity is not working". In this regard, however, our analysis offers a second important insight: if policy makers and the electorate show sufficient resolve, carrying through with austerity will eventually be rewarded by better financing conditions. Austerity pays off in the long run.

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A Appendix

A.1 Details on the construction of the default premium

In this subsection, we provide additional information on the construction of default premia and data sources.

A.1.1 EMBI spreads

The J.P. Morgan EMBI is an emerging market debt benchmark that includes "U.S.-dollar-denominated Brady bonds, Eurobonds, traded loans, and local market debt instruments issued by sovereign and quasi-sovereign entities" (JP Morgan, 1999). For our purposes, it is important to note that debt instruments must have at least 2.5 years of maturity left for inclusion and remain in the index until 12 months before maturity. This implies that the maturity of the EMBI does not necessarily stay constant over time as the maturity of the underlying debt portfolio may change. The EMBI spread "corresponds to the weighted average of these securities' yield difference to the US Treasury securities with similar maturity, considered risk free. This risk premium is called in the market as the spread over Treasury of this portfolio" (Banco Central do Brasil, 2014). Inclusion of a bond into the EMBI requires a minimum bond issue size of $500 million. This ensures that the liquidity premium compared to U.S. bonds is not too large.[35]

The data is retrieved from Datastream. The mnemonic is JPMG followed by a three letter country identifier. We rely on stripped spreads (Datastream Mnemonic: SSPRD), which "strip" out collateral and guarantees from the calculation. For example, JPMGARG(SSPRD) is the mnemonic for the Argentinean EMBI spread.

A.1.2 CDS spreads data

CDS spreads are from Datastream and spliced from two sources. Until 2010Q3, Datastream provides CDS spreads from Credit Market Analysis Limited (CMA), while Thomson Reuters, starting in 2008 provides CDS for an increasing number of issuers.[36] The contract type we choose is five years of maturity with complete restructuring (CR). The CMA CDS spreads are typically denominated in dollar, while the Thomson Reuters CDS spreads are often available in euro and dollar. Despite CDS spreads being theoretically unit free as they are measured in basis points, the choice of denomination currency choice can be relevant for sovereign entities. The reason is that, e.g., being reimbursed in U.S. dollar when Germany defaults may provide an insurance against exchange rate risk. (for more on this and CDS contracts in general, see, e.g., Buchholz and Tonzer, 2013; Fontana and

[35] For more information on the EMBI see JP Morgan (1999). Banco Central do Brasil (2014) provides a very accessible general introduction to the EMBI.

[36] Additional information on the distinction and the how to match the two series can be found at `http://extranet.datastream.com/data/CDS/`.

Scheicher, 2010). To exclude an exchange rate risk premium, we use Thomson Reuters CDS spreads in U.S. dollar for all non-EMU countries and Thomson Reuters Euro CDS spreads in euro for euro area members after EMU accession. Unfortunately, for early time periods, the currency-specific Thomson Reuters CDS spreads are not always available. In this case, we rely on the CMA CDS spreads.

A.1.3 Spread decomposition

In the main text, we use the difference between nominal yields on foreign-currency bonds and a risk-free reference bond to measure the default premium. We elaborate on this in the following.

For most practical purposes, the nominal yield to maturity of a bond, r_t^{nom} can be decomposed as

$$r_t^{nom} = r_t^{real,riskfree} + E_t\left(\pi_{t+1}\right) + RP_t^{Infl} + E_t\left(\delta_{t+1}\right) + RP_t^{default} + RP_t^{term} + RP_t^{liqu} + \varepsilon_t, \quad \text{(A.1)}$$

where $r_t^{real,riskfree}$ is the real risk-free interest rate, $E_t\left(\pi_{t+1}\right)$ is the compensation for expected inflation, RP_t^{Infl} denotes the premium for inflation risk, and RP_t^{term} the term premium.[37] We are mostly interested in the next two components that we subsume under the heading "default premium": the compensation for expected default $E_t\left(\delta_{t+1}\right)$ and the default risk premium $RP_t^{default}$. The term RP_t^{liqu} captures liquidity risk premia, while ε_t captures other (higher order) terms. In order to isolate the terms of interest to us, we compute the yield spread between foreign-currency bonds and a default-risk free reference bond/bond index of a similar maturity. Under integrated financial markets, its yield, $r_t^{*,nom}$, will be given by

$$r_t^{*,nom} = r_t^{real,riskfree} + E_t\left(\pi_{t+1}\right) + RP_t^{Infl} + RP_t^{term} + RP_t^{*,liqu} + \varepsilon_t^* . \quad \text{(A.2)}$$

The default-related terms are zero. The real risk-free interest rate, the inflation premium, and the term premium should be the same as in Equation (A.1), as we consider a bond denominated in the same currency and with the same maturity.[38] A yield spread computed this way will thus only contain the default-related premium and the difference of the liquidity risk premium as well as higher order terms. Unfortunately, it is not easy to isolate the difference in liquidity premia. However, we are quite confident that liquidity is not driving our results for three reasons. First, markets for government bonds are typically quite liquid so that any liquidity premium should be small. Second, the risk premium consist of the price of risk times the quantity of risk. With integrated financial markets, the price of risk tends to be a common factor that will be accounted for by our time-fixed

[37] This is a second order effect arising from the covariance of returns with the stochastic discount factor. It is absent if all investors are risk neutral.

[38] Regarding the term premium, it is actually the duration of expected cash flows that matters. This might introduce small differences of the term premium (see Broner et al., 2013).

effects, leaving only the quantity component of liquidity risk as a confounding factor (see also the discussion in Section A.2.). Finally, we find that our main results also obtain for a sample of developed economies where markets are very liquid. Results also hold up if we drop observations after the beginning of the recent financial crisis—a period when liquidity dried up considerably.

A.2 Price of risk and quantity of risk

Our measure of the default premium reflects the quantity of risk times the price of risk. The price of risk may be time-varying with global risk aversion (see, e.g., Bekaert et al., 2013). However, this should not be a problem in our setup as the price of risk-component should be global and is thus captured by time-fixed effects. This is equivalent to including the VIX as a control. However, our fiscal stress indicator is also based on default premia and thus depends on the price of risk as well. Thus, while the cross-section of our fiscal stress indicator is unaffected by the price of risk, the time series dimension may be affected as the price of risk will be simultaneously high for all countries at a particular point in time. However, results are robust to dropping the Great Recession period from our sample—a period when price of risk spiked.

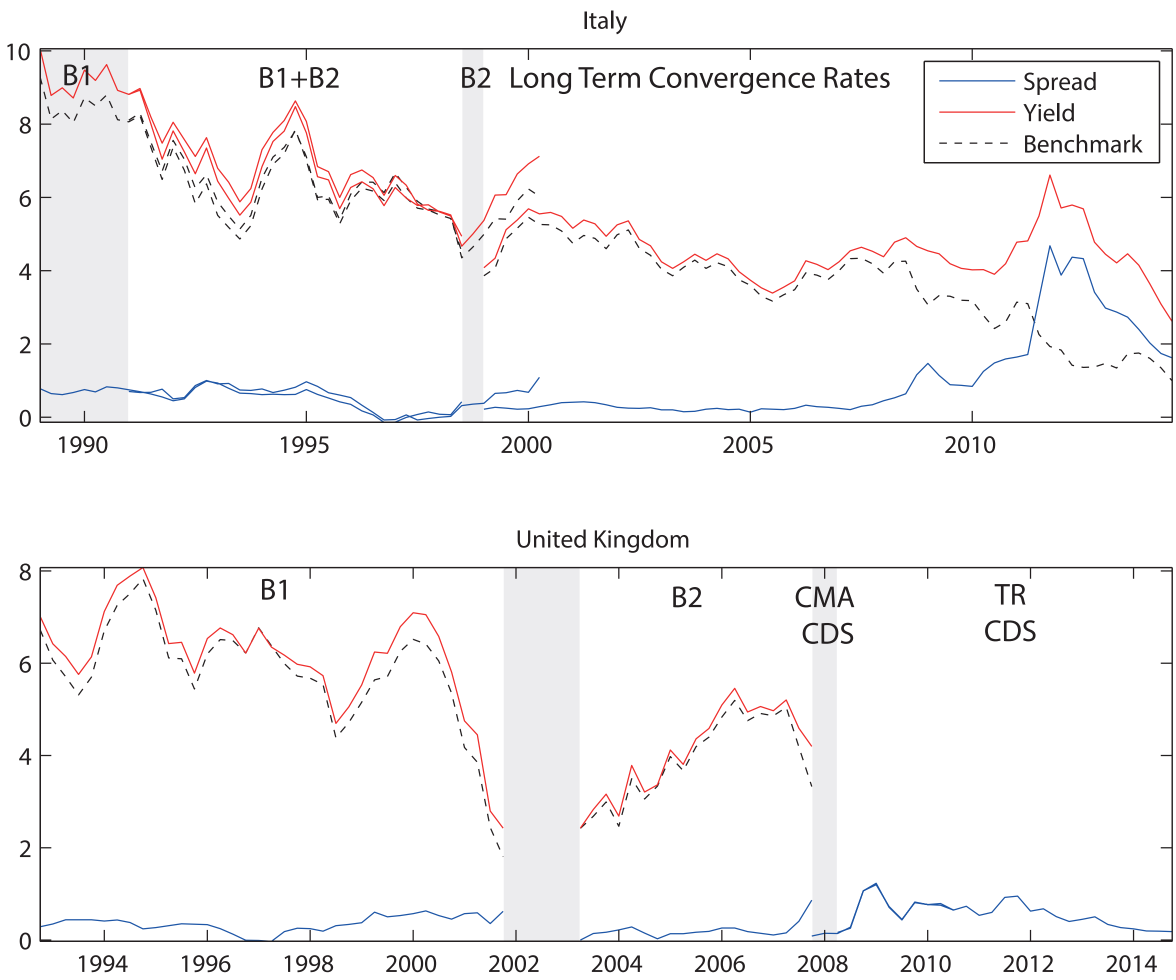


Figure A.1: Construction of sovereign yield spread: Italy and United Kingdom.

A.3 Construction of data set: example

To illustrate the construction of our data set, Figure A.1 provides two examples, namely data for Italy (top) and the United Kingdom (bottom). Until 1991 only one Italian foreign currency-bond is available. Starting in 1992, we obtain a second bond and compute the yield spread as the average over those bonds. When the first bond matures in 1997, we are left with one bond until 1999. From that point on, we use the long-term convergence bond yields provided by the ECB. For the United Kingdom, we have two different bonds available to cover the early part of the sample, with missing values in between. From 2007 on, we rely on CMA CDS spreads, while in 2008 the Thomson Reuters CDS spreads become available, which are used for the rest of the sample.[39]

[39]For details, see Appendix A.1.2.

B Additional tables and figures

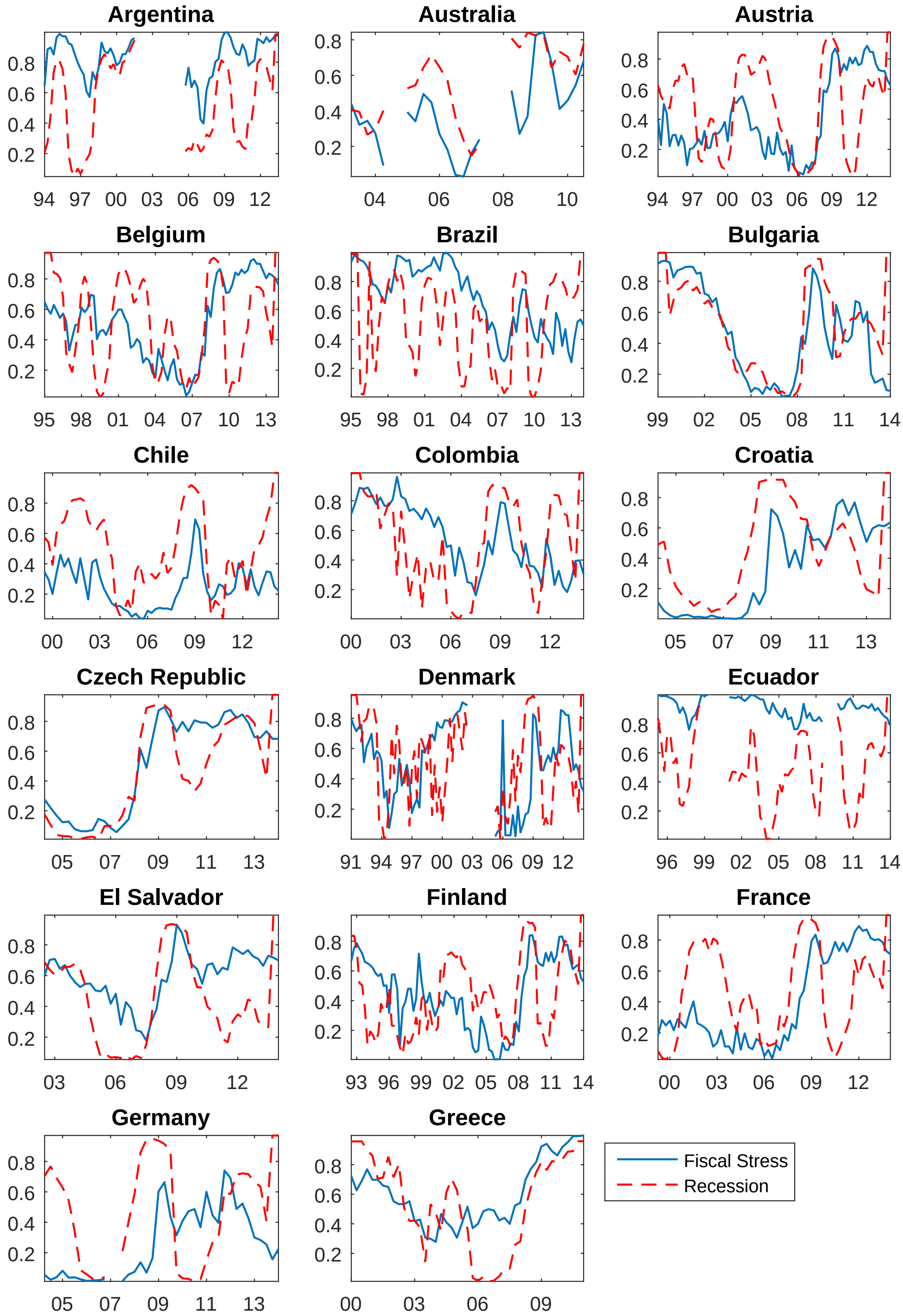


Figure A.2: Values of empirical CDF (Country group-specific) for lagged default premia and smoothed output gaps.

Table A.1: Descriptive statistics: indicator functions

Country	$mean\,(F^{stress})$	$mean\,(F^{recess})$	$corr\,(F^{stress}, F^{recess})$
Argentina	0.82	0.53	0.60
Australia	0.39	0.56	0.69
Austria	0.40	0.48	0.34
Belgium	0.51	0.51	0.40
Brazil	0.68	0.53	0.19
Bulgaria	0.46	0.52	0.70
Chile	0.25	0.51	0.63
Colombia	0.54	0.54	0.30
Croatia	0.31	0.45	0.53
Czech Republic	0.50	0.46	0.85
Denmark	0.50	0.52	0.33
Ecuador	0.91	0.50	0.05
El Salvador	0.59	0.45	0.55
Finland	0.45	0.44	0.31
France	0.40	0.47	0.27
Germany	0.27	0.49	0.11
Greece	0.58	0.55	0.70
Hungary	0.28	0.44	0.43
Ireland	0.49	0.51	0.12
Italy	0.60	0.51	0.25
Latvia	0.76	0.51	0.52
Lithuania	0.73	0.46	0.53
Malaysia	0.27	0.51	0.53
Mexico	0.55	0.49	0.11
Netherlands	0.37	0.50	0.37
Peru	0.56	0.56	0.30
Poland	0.32	0.53	0.44
Portugal	0.49	0.52	0.13
Slovakia	0.54	0.44	0.80
Slovenia	0.58	0.45	0.64
South Africa	0.41	0.48	0.56
Spain	0.48	0.51	0.54
Sweden	0.47	0.52	0.04
Thailand	0.20	0.52	0.40
Turkey	0.66	0.51	0.22
United Kingdom	0.54	0.44	0.34
United States	0.39	0.55	-0.17
Uruguay	0.58	0.51	-0.05

Notes: F^{stress} denotes the values of the country group-specific empirical CDF of the lagged default premium; F^{recess} denotes the empirical CDF of the smoothed output gap, computed as the z-scored deviation of the 5 quarter moving average of the output growth rate from its HP-filtered trend ($\lambda = 160{,}000$). First column: average value of the fiscal stress indicator for the respective country. Second column: average value of the recession indicator for the respective country. Last column: correlation between the two indicators.

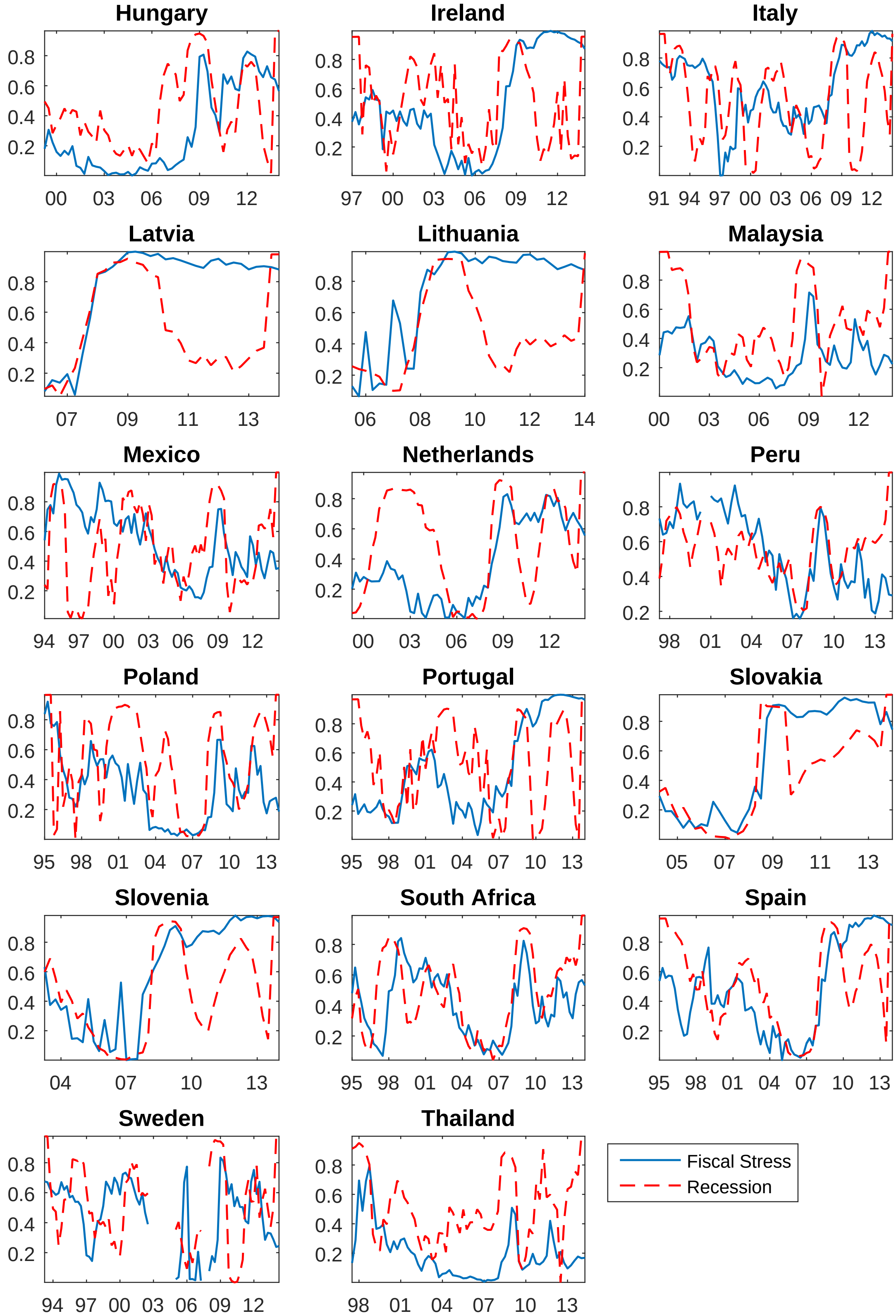


Figure A.3: Values of empirical CDF (Country group-specific) for lagged default premia and smoothed output gaps.

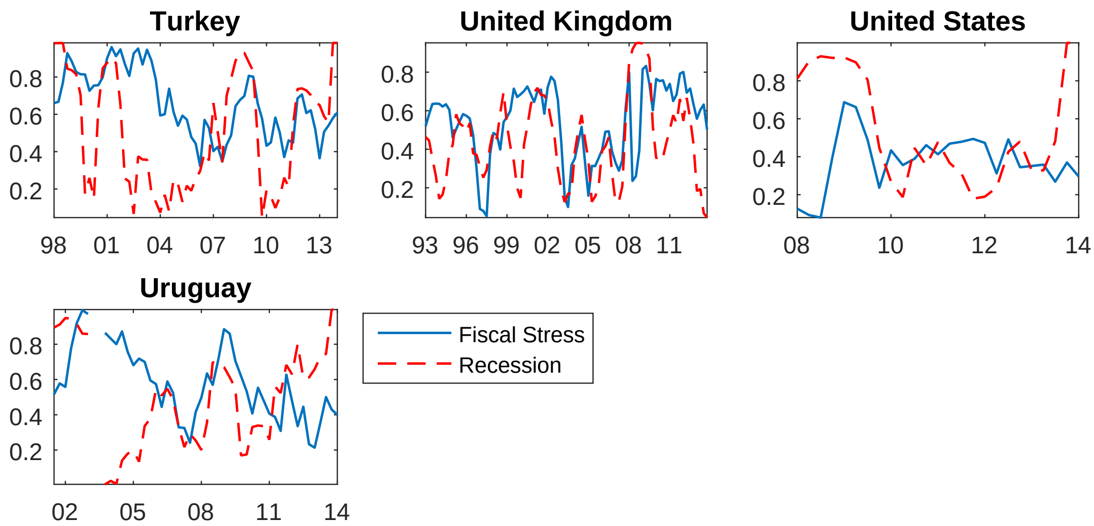
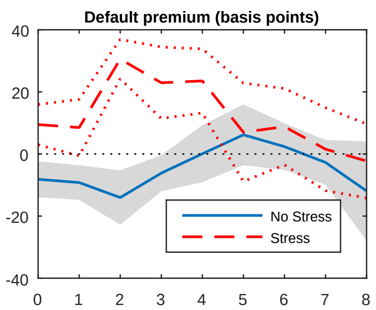
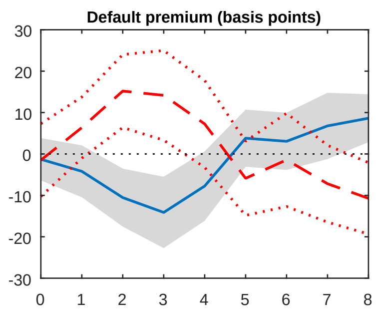


Figure A.4: Values of empirical CDF (Country group-specific) for lagged default premia and smoothed output gaps.

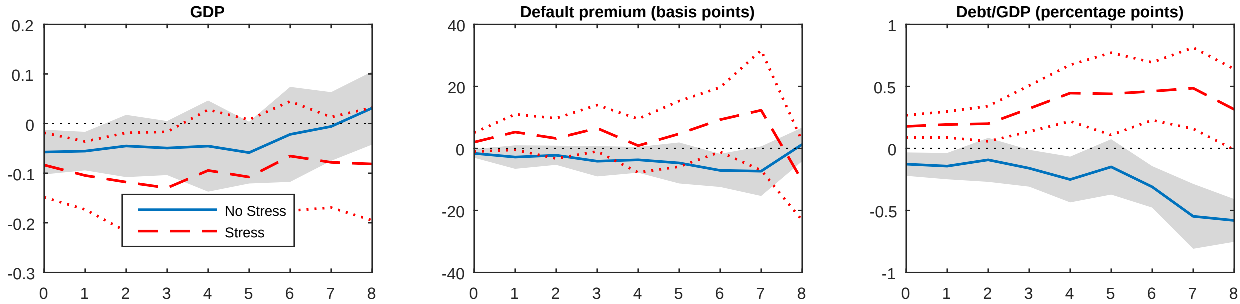


(a) Forecast error of government consumption

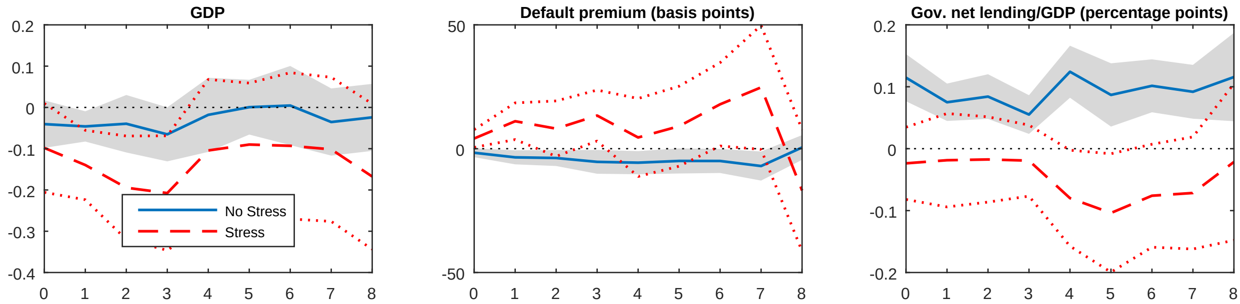


(b) Actual government consumption

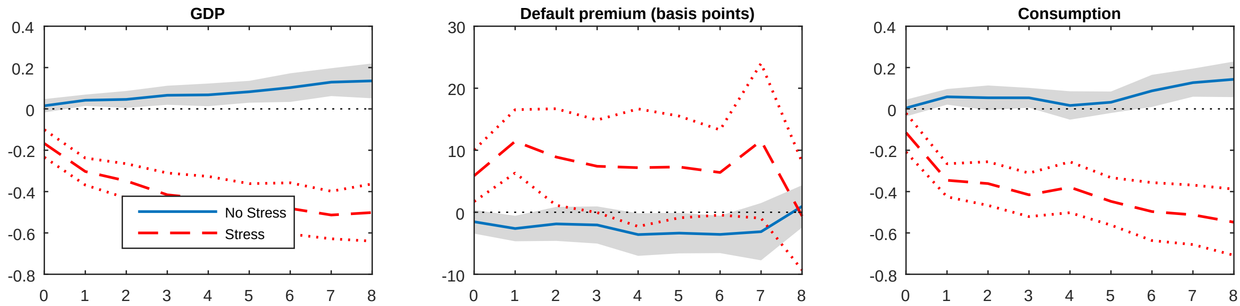
Figure A.5: Dynamic response of default premium to cut of government consumption by 1 percent: semi-annual observations. Results based on local projections conditional on the absence of fiscal stress (solid lines) and on the presence of stress (dashed lines). Panel (a): responses based on forecast error of government consumption, Panel (b): responses based on government consumption (OECD sample of semi-annual observations).



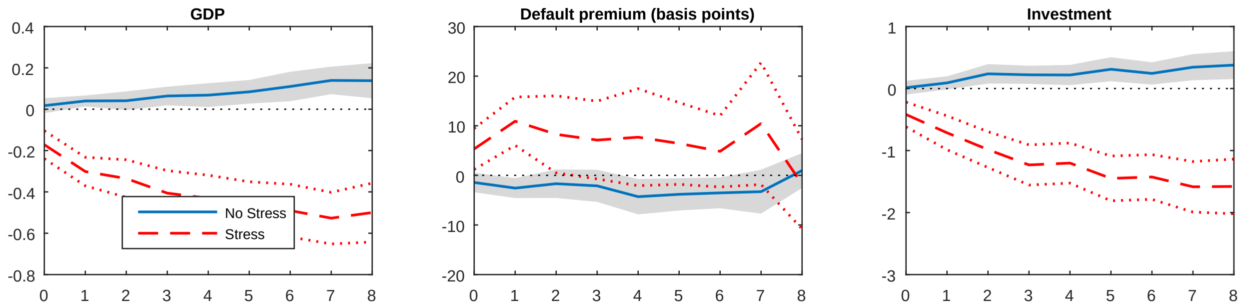
(a) Including debt/GDP ratio



(b) Including government net lending/GDP ratio

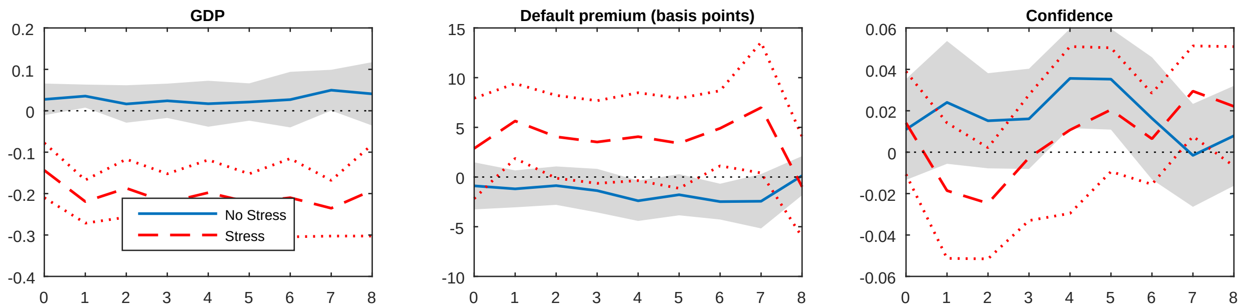


(c) Including private consumption

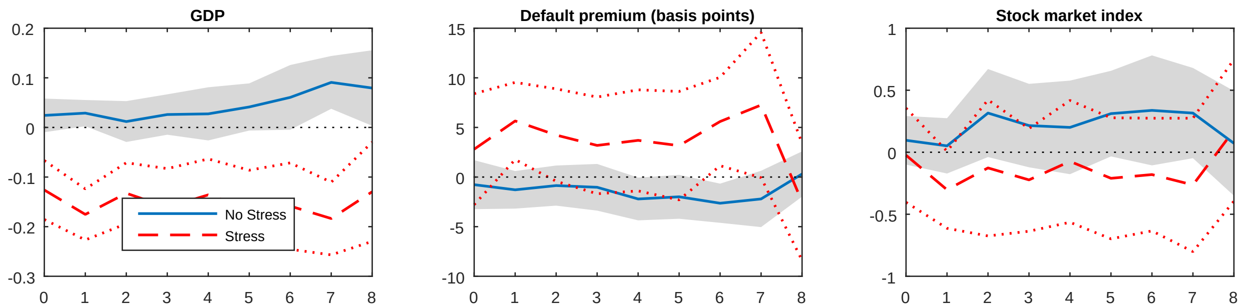


(d) Including private investment

Figure A.6: Dynamic response to cut of government consumption by 1 percent in the four-variable model. Notes: see Figure 4.

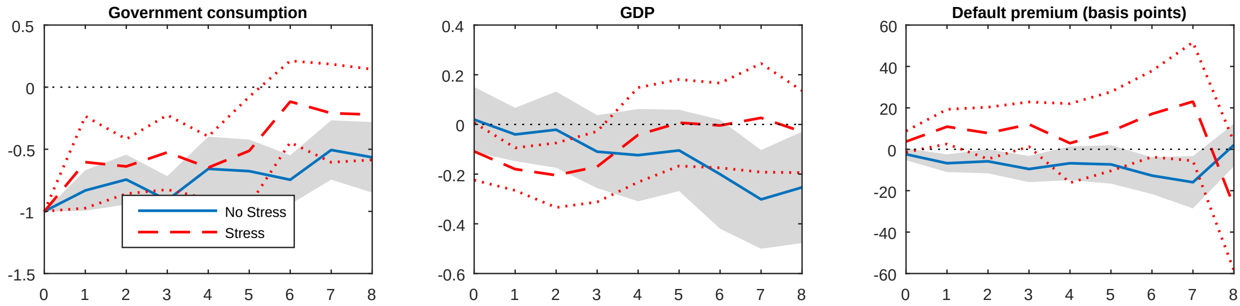


(a) Including confidence

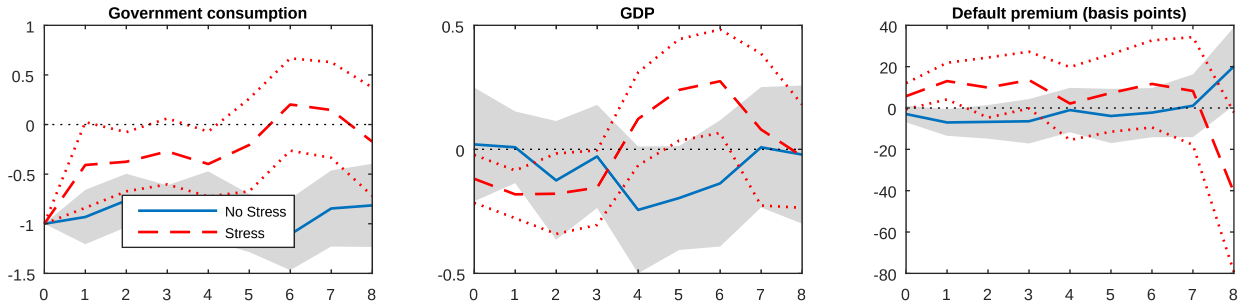


(b) Including stock market returns

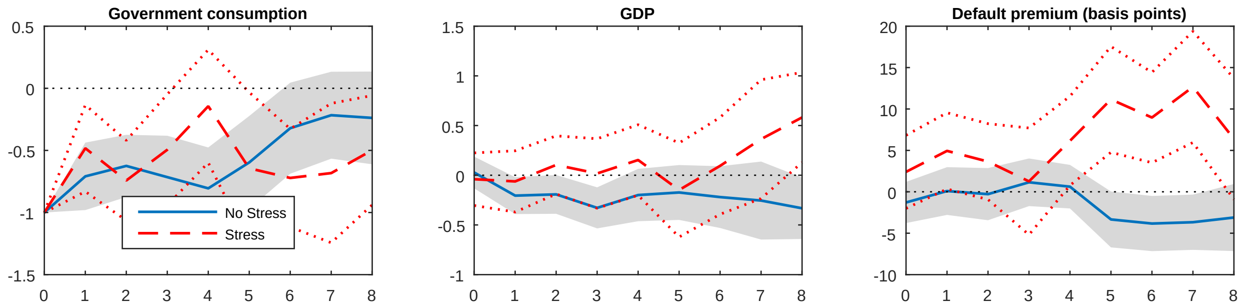
Figure A.7: Dynamic response to cut of government consumption by 1 percent in the four-variable model. Notes: see Figure 4.



(a) Euro area

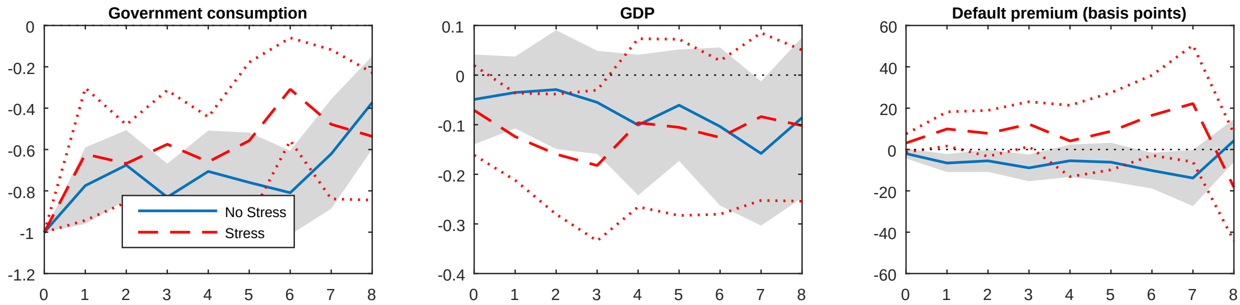


(b) Euro area: crisis countries

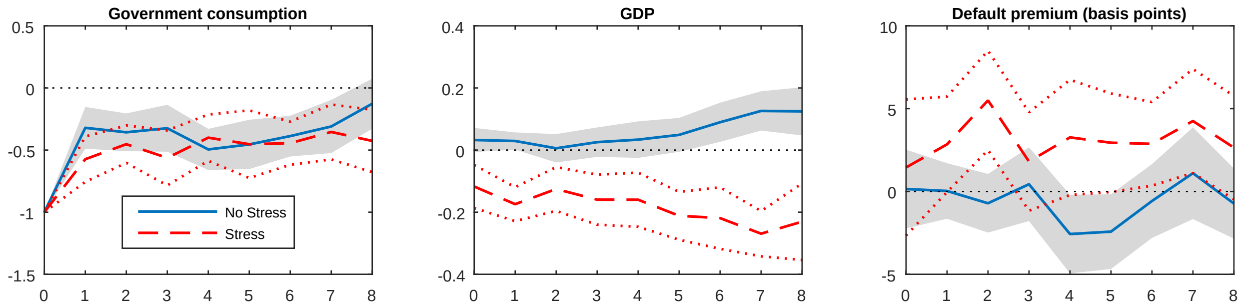


(c) Euro area: non-crisis countries

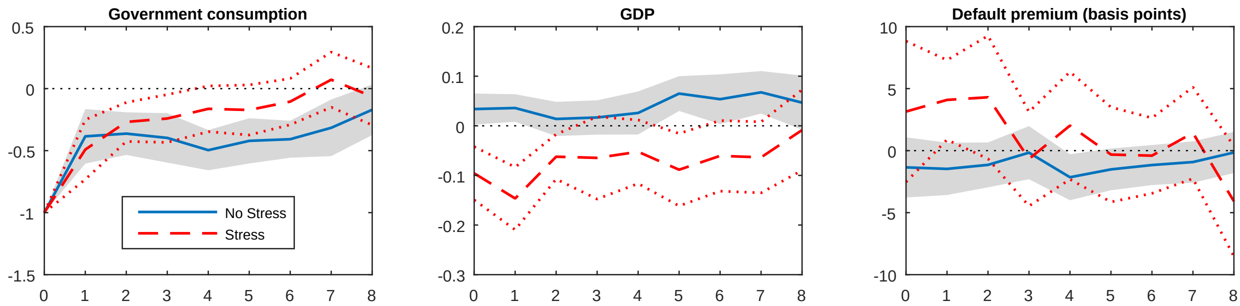
Figure A.8: Dynamic response to cut of government consumption by 1 percent: euro area samples. Notes: see Figure 4.



(a) Developed economies

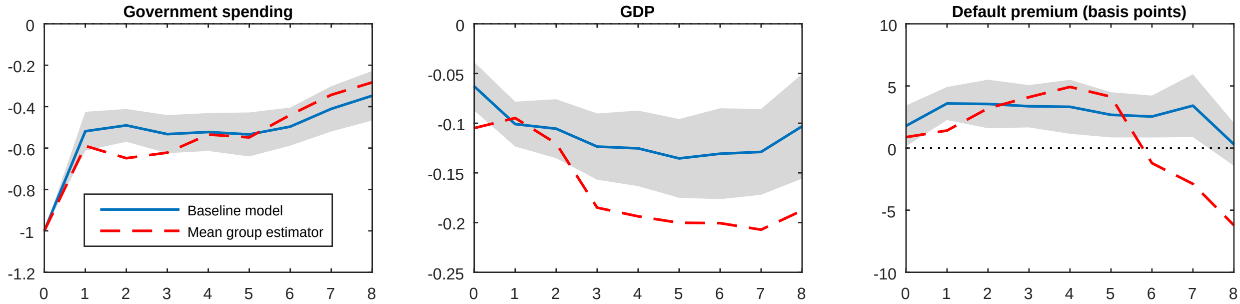


(b) Emerging economies

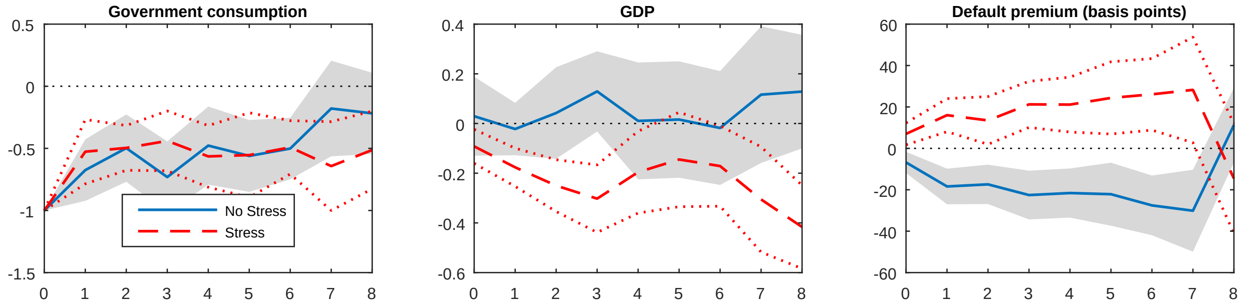


(c) Excluding Great Recession: conditional on fiscal stress

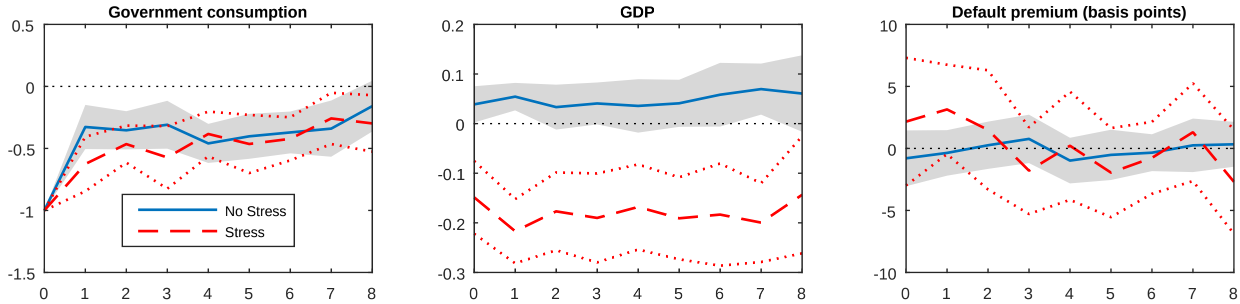
Figure A.9: Dynamic response to cut of government consumption by 1 percent: panels (a) and (b) show results for local projections on developed and emerging economies separately; panel (c) shows results for when the Great Recession period is dropped from the sample. Notes: see Figure 4.



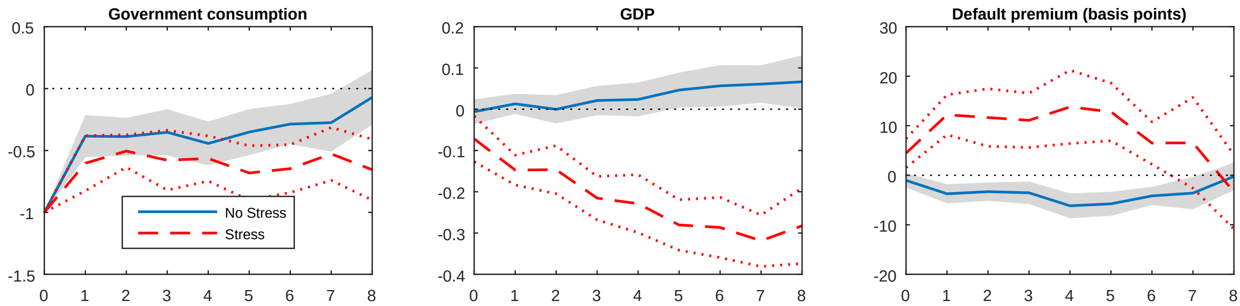
(a) Unconditional model: baseline vs. mean group estimator



(b) Monetary union or dollarization

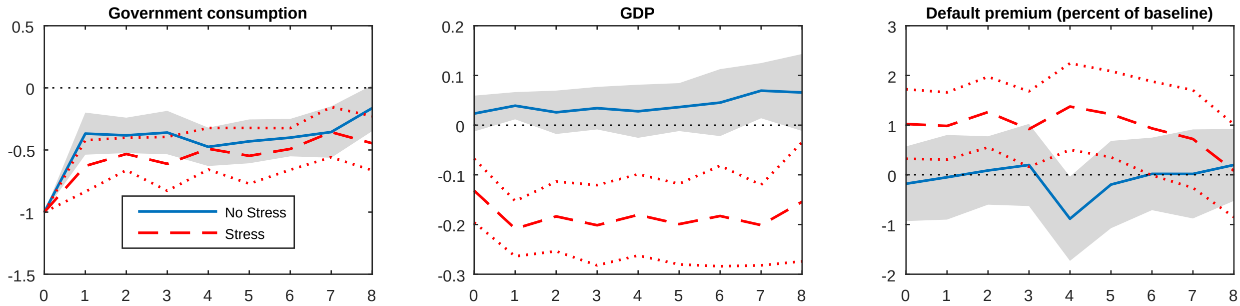


(c) Countries with their own legal tender

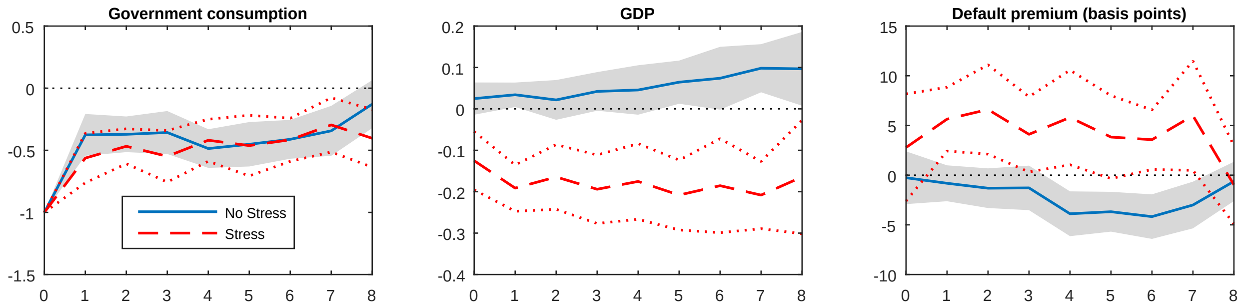


(d) IMF program countries excluded

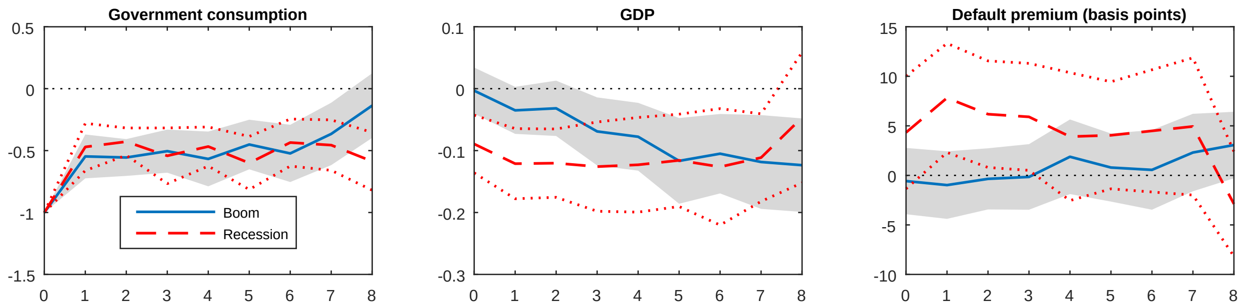
Figure A.10: Dynamic response to cut of government consumption by 1 percent. Panel (a): comparison of unconditional baseline estimates with those obtained using mean group estimator. Panel (b): sample includes only country-quarter observations for countries which are members of a monetary union or de jure dollarized. Panel (c): only country-quarter observations when countries have their own legal tender. Panel (d): only country-quarter observations w/o IMF program. Notes: see Figure 4.



(a) Default premia measured in logs



(b) Conservative sample for government consumption



(c) Boom vs. recession

Figure A.11: Dynamic response to cut of government consumption by 1 percent. Panel (a): default premia measured in logs. Panel (b): conservative sample where we could confirm that government spending data was derived from direct sources. Panel (d): conditioning on booms and recessions (output gap used as indicator variable). Notes: see Figure 4.